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1 The Almost Ideal Demand System: A Comparison and Application to Food Groups

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40 The Food and Agricultural Policy Simulator: An Addendum

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In This Issue

This issue emphasizes the ties between agricultural economists and other economists. Both research articles apply modeling techniques from the general literature to processes of particular concern to agricultural economists—food purchases and corn production.

Blanciforti and Green apply the almost ideal demand system (AIDS) from the general literature to four categories of food purchases. This new demand system has several advantages over the well-known linear expenditure system (LES), which they employ for comparison. The AIDS is indirectly nonadditive, allowing consumption of one good to affect the marginal utility of another, whereas the LES is directly additive, implying independent marginal utilities. Furthermore, the AIDS possesses the property that income elasticities become more inelastic for necessities (for example, food items) as their budget shares decrease. The reverse is true for the LES. Both the AIDS and LES classify meats as relative luxuries (food expenditure elasticities greater than 1) and cereals and bakery goods as relative necessities. However, the classification of fruits and vegetables switches from a relative necessity under LES to a relative luxury under AIDS, whereas the classification for miscellaneous foods switches in the opposite direction.

In the second article, Larson illustrates the potential pitfalls of relying on R^2's and t-values when a forecasting model is developed. He presents a structural model, similar to some models in use, of the corn market with very "good" values of the commonly employed statistics. He then demonstrates the model's poor forecasting performance and goes on to transform the model so that the values of the standard statistics are "poorer," but the forecasting performance is better. Other potential forecasting models—ARIMA, vector autoregression, and open autoregression—are estimated for comparison, but they do not forecast as well as the transformed structural model.

The Research Review section summarizes an important topic in macroeconomics and monetary theory. With interest expenses rising from 8 percent of farm production expenses in 1975 to 16 percent in 1982, monetary theory and policy have captured more of the agricultural economist's attention. The effects of the economic slowdown of the past few years have further emphasized the importance of the macroeconomy to agriculture.

A large part of professional economic discourse in monetary theory and macroeconomics currently focuses on the merits and demerits of the rational expectations literature and its corollary: that discretionary monetary policy cannot permanently alter the real variables in the economy—the level of economic activity, the level of unemployment, and inflation-adjusted interest rates. Conway and Barth summarize the important literature in this debate, including articles that change the basic model to imply policy effectiveness, generally produced by long adjustment lags. They conclude that the lasting contributions of this literature may be to encourage closer attention to modeling expectations and to stimulate the responses of economic agents to policy actions.

The rational expectations hypothesis would imply small fluctuations in expected real interest rates. However, high real interest rates have persisted for some time. Sundell summarizes the literature on the determination of real interest rates, a literature which partially depends on the rational expectations literature. He concludes that real interest rates have been quite variable because of slow adjustments of wages, prices, and output to changes in the nominal money supply.

Lorna Aldrich

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The Almost Ideal Demand System: A Comparison and Application to Food Groups

By Laura Blanciforti and Richard Green*

Abstract

This article presents estimates of the almost ideal demand system (AIDS) for four food groups and compares these estimates with the AIDS' own linear approximate version and the linear expenditure system. The AIDS is indirectly nonadditive and has several desirable properties, making it a viable demand system for analyzing food commodities. Its linear approximate version is a good first-order approximation to the complete system and is easy to estimate.

Keywords

Demand, systems, food groups

Introduction[1]

Demand theory is concerned with the allocation of total expenditures among goods and services, given prices and consumer income. The focus on total expenditures, rather than on expenditures for a single commodity, makes it possible to examine interdependencies among commodities. Although single-equation demand functions have the advantage of modeling a commodity in isolation and of allowing far more flexibility in accounting for explanatory variables and specification of functional forms, the demand system approach accounts for interdependencies among commodities, includes theoretical restrictions, is often derived from a utility maximization process,[2] and describes the allocation of expenditures among a complete set of consumption categories that sum to total expenditures.

Given the parameters for a complete demand system, a researcher could simulate, for example, the effect of a sharp increase in housing or energy prices on food expenditures. If such simulations are to be used by policymakers, however, they must emerge from systems with both plausible assumptions and results. Economic Research Service economists have used several complete demand systems to examine food expenditures, but earlier stages in the development of complete systems have required them to use systems with some implausible assumptions. This article examines food expenditures with a system which is more realistic than those used earlier and which, in its linear approximate form, is easy to estimate.

The complete system approach was pioneered by Stone (*20*), who developed a system consistent with the assumptions of neoclassical demand theory and was able to estimate it with data for Great Britain by combining commodities into manageable groups. In an interesting application of the linear expenditure system (LES), Stone assessed the effect of rationing in Great Britain by simulating desired expenditures at prices that existed under rationing. However, Stone's system restricted the nature of the relationship of commodities by assuming that the underlying preference ordering was additive—that is, that the marginal utility provided by the consumption of one commodity was independent of the consumption of other commodities. The results were that all goods were substitutes and inferior goods were excluded.

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[1] For a more detailed treatment of complete demand systems, see (*2*). Italicized numbers in parentheses refer to items in the References at the end of this article.

[2] Exceptions include the double logarithmic and other *ad hoc* systems.

Strotz (*22*) extended the idea of exhaustive expenditures to stages. In the first stage, the consumer is assumed to allocate expenditures to broad groups of commodities; then, in the second stage, the consumer is assumed to allocate expenditures within each of the broad groups to smaller groups. This process can continue, but for most empirical analyses has been limited to two stages requiring the condition of weak separability—that is, the conditional ordering of goods based on the independence of marginal utilities of goods within one group from consumption of goods in other groups.[3]

Deaton and Muellbauer (*10*) recently extended empirical research on demand systems by developing and estimating the almost ideal demand system (AIDS). The name stems from the properties associated with their system. Deaton and Muellbauer (*10*, p. 312) list the following advantages of their system: (1) it gives an arbitrary first-order approximation to any demand system; (2) it satisfies the axioms of choice exactly; (3) it aggregates perfectly over consumers; (4) it has a functional form which is consistent with previous household budget data; (5) it is simple to estimate in its linear approximate form; and (6) it can be used to test for homogeneity and symmetry. In addition, although Deaton and Muellbauer do not explicitly mention it, the AIDS is indirectly nonadditive, allowing consumption of one good to affect the marginal utility of another good; whereas, the linear expenditure system is directly additive, implying independent marginal utilities. Thus, the AIDS, in addition to the listed desirable properties, does not impose the severe substitution limitations implied by additive demand models such as the LES.

Our purpose here is to report results obtained from applying the new AIDS to a four-food (second-stage) commodity classification. Thus, assuming weak separability, we can focus on the allocation of food expenditures among this particular set of nondurable goods. This subsystem demand approach allows us to compare substitution possibilities among these food types. These estimates account for restrictions imposed by theoretic demand formulations. Although the system presented here could benefit from more disaggregation, it attempts to estimate a theoretically plausible, complete demand system for a major commodity group and is a first step toward understanding the relationship among commodities. In addition, we make comparisons with a simplified linear approximation of the AIDS and with the LES. The latter system, while admittedly somewhat inappropriate for use with such a highly refined food grouping, serves as a benchmark for evaluating the results from the more viable AIDS.

Based on U.S. annual time series data for 1948-78, the findings of our analysis indicate that many commodities classified as luxuries in the LES because their income elasticities are greater than 1, are classified as necessities in the AIDS as their income elasticities are less than 1. The less restrictive AIDS does not reflect an approximate proportional relationship between income and price elasticities as is often found when one uses the LES (for example, see (*9*)).

Besides the properties of the AIDS described by Deaton and Muellbauer (*10*), we show the AIDS possesses the property that income elasticities become more inelastic for necessities (for example, food items) as their budget shares decrease. The reverse is true for the LES. Thus, the AIDS is an attractive system for analyzing the demand for food commodities. Excluding its linear-approximate version, one disadvantage is that it requires a large number of parameters to be estimated.

Models

We chose the two demand systems, the LES and the AIDS, based on theoretical and empirical considerations. Both these demand systems are complete, theoretically plausible systems and satisfy the properties of demand systems. However, the LES results are reported primarily to help us evaluate the results obtained from the AIDS. We briefly describe the LES and give an indepth account of the AIDS because it is less well known than the LES.

Linear Expenditure System (LES)

The LES, which can be derived from the Stone-Geary utility function, in budget share form, is given by:

$$w_i = p_i\mu_i/Y + \theta_i(1 - \sum_k p_k\mu_k/Y) \text{ for } i,k = 1,\ldots,n \qquad (1)$$

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[3] See (*11*, p. 124) and (*6*, pp. 287-88).

where the w_i's are budget shares, the p_i's are prices, the μ_i's are interpreted as minimum required subsistence quantities, the θ_i's are marginal budget shares, and Y is total expenditure (income). It can be shown (*12*) that the LES globally satisfies the adding up, homogeneity, and symmetry restrictions. The LES is also described as an additive system because it is derived from an additive utility function.[4]

To estimate the LES, we impose the condition that the marginal budget shares aggregate to 1 and impose cross-equation restrictions which are implied by theory. If the quantities consumed are positive and greater than their minimum subsistence levels and the marginal budget shares are valued between zero and 1, the elasticities will have their typical pattern—that is, positive income elasticities, exclusion of inferior goods, and negative own-price elasticities. Because of its additive form, the LES has been shown by Deaton (*9*) to imply an approximate proportional relationship between income elasticities and own-price elasticities, commonly referred to as the Pigou relationship.

In addition to being a theoretically plausible demand system (that is, derived from a utility maximization process), having an intuitive economic interpretation, and being relatively easy to estimate, the LES has performed well in terms of goodness of fit, prediction, and so forth (*14, 15*) in comparison with nonadditive systems.

Almost Ideal Demand System (AIDS)[5]

The new demand system—AIDS—developed by Deaton and Muellbauer (*10*) builds upon a model by Working (*26*) and Leser (*16*). Their model expresses the ith budget share, w_i, as a function of log Y, that is:

$$w_i = \alpha_i + \beta_i \log Y \qquad (2)$$

where w_i and Y are defined as above. The Working-Leser model was extended by Deaton and Muellbauer to include the effect of prices. The resultant demand system for the AIDS was derived, by use of duality concepts, from a particular cost or expenditure function defined as the minimum expenditure necessary to attain a specific level of utility at given prices. Thus, it is also a theoretically plausible demand system. Consider the cost function (*10*, p. 313):

$$\log C(U,p) = \alpha_o + \sum_k \alpha_k \log p_k + 1/2 \sum_k \sum_j \gamma^*_{kj} \log p_k \log p_j + U\beta_o \prod_k p_k^{\beta_k} \qquad (3)$$

where C denotes the cost function, U represents the unobservable utility parameter, β_o is a nonestimable cost parameter, p_k's are prices, and α_k, γ^*_{kj}, and β_k are parameters to be estimated. Deaton and Muellbauer chose the particular form in equation (3) to allow the cost function to be flexible, to represent preferences via the cost function that permit exact aggregation over consumers, and to obtain a system of demand functions with desirable properties. By applying Shepard's Lemma, that is, by differentiating equation (3) with respect to prices, they obtain the Hicksian, or compensated demand functions. Mathematically:

$$\frac{\partial C(U,p)}{\partial p_i} = q_i(U,p) = q_i \qquad (4)$$

By multiplying both sides by $p_i/C(U,p)$, equation (4) becomes:

$$\frac{\partial \log C(U,p)}{\partial \log p_i} = \frac{\partial C(U,p)}{\partial p_i} \bullet \frac{p_i}{C(U,p)} = \frac{p_i q_i(U,p)}{C(U,p)} \qquad (5)$$
$$= w_i(U,p)$$

For the cost function given by equation (3), equation (5) becomes:

$$w_i = \alpha_i + \sum_j \gamma_{ij} \log p_j + \beta_i U \beta_o \prod_k p_k^{\beta_k} \qquad (6)$$

where $\gamma_{ij} = 1/2(\gamma^*_{ij} + \gamma^*_{ji})$.

Because $Y = C(U,p)$ in equilibrium, by substituting Y for C in equation (3), then by solving for U in terms of p and Y, and finally by substituting this expression into equation (6), we obtain the AIDS in budget share form:

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[4] A utility function is additive if there is a differentiable function F, $F' > 0$, and n functions $f_i(q_i)$, so that $F[f(q_1,\ldots,q_n)] = \sum_i f_i(q_i)$. The Stone-Geary utility function $U(q) = \sum_i \theta_i \log(q_i - \mu_i)$ satisfies this condition. See (*19*, pp. 57-58).

[5] As a point of interest, the first difference form of the AIDS is similar to the Rotterdam demand system. The results from the estimation of a Rotterdam system were presented in this journal by Mann (*17*).

$$w_i = \alpha_i + \sum_j \gamma_{ij} \log p_j + \beta_i \log(Y/P), \qquad (7)$$

$$\text{for } i,ji = 1,\ldots,n$$

where P is a price index defined by:

$$\log P = \alpha_o + \sum_k \alpha_k \log p_k + 1/2 \sum_k \sum_j \gamma^*_{kj} \log p_k \log p_j \text{[6]} (8)$$

Deaton and Muellbauer (*10*) utilize Stone's (*21*) index ($\log P^* = \sum_k w_k \log p_k$), where $P \cong \xi P^*$; that is, P is assumed to be approximately proportional to P^*, and they apply ordinary-least-squares (OLS) estimation. Thus, equation (7) is redefined as:

$$w_i = \alpha_i^* + \sum_j \gamma_{ij} \log p_j + \beta_i \log(Y/P^*) \qquad (9)$$

where $\alpha_i^* = \alpha_i - \beta_i \log \xi$. This equation will be referred to as the linear approximate/almost ideal demand system (LA/AIDS) and is often a good first-order approximation to the complete AIDS system, equation (7).

In this form, with P as a price index, the coefficients are easily interpreted. The ith budget share is expressed in terms of prices and real income or expenditures, Y/P. The α_i is the intercept and represents the average budget share when all logarithmic prices and real expenditures are equal to 1. The γ_{ij} is equivalent to the change in the ith budget share with respect to a percentage change in the jth price with real expenditures or income held constant; that is, $\gamma_{ij} = \partial w_i / \partial \log p_j$. The β_i represents the change in the ith budget share with respect to a percentage change in real income or expenditures with prices held constant; that is, $\beta_i = \partial w_i / \partial \log(Y/P)$.

The demand properties (commonly known as adding up, homogeneity, and Slutsky symmetry) can be shown to be satisfied for the AIDS. First, for adding up, the budget shares sum to 1 if $\sum_i \alpha_i = 1$, $\sum_i \gamma_{ij} = 0$, and $\sum_i \beta_i = 0$. Second, the homogeneity condition holds if $\sum_j \gamma_{ij} = 0$. And, finally, the symmetry restriction holds if $\gamma_{ij} = \gamma_{ji}$. Deaton and Muellbauer rejected the latter conditions, and we test them in this analysis.

In the complete AIDS, equation (7), notice that there are $2n + n^2$ parameters to be estimated—n α_i's, n β_i's, and n^2 γ_{ij}'s. The number of restrictions just mentioned totals $(n^2 + n + 4)/2$.[7] These restrictions reduce the number of free, unknown structural parameters to $(n^2 + 3n - 4)/2$. In any case, many parameters must be estimated in the AIDS. As the number of commodities, n, increases, the total number of parameters to be estimated multiplies, and this could result in estimation problems.[8] With this in mind, we chose four commodities for our analysis.

With reference to the LES, there are only 2n structural parameters—n θ_i's and n μ_i's; and with one restriction, the $\sum_i \theta_i = 1$, there are $2n - 1$ unknown parameters.

Both equations (1) and (7), the LES and AIDS in budget share form, are nonlinear; and full information maximum likelihood (FIML) procedures can be used for maximum efficiency in estimation. Equation (9), the LA/AIDS, is linear because the $\log P^*$ term is an exogenous approximation, is estimated by OLS procedures, and is used to examine homogeneity.[9] Homogeneity is tested by imposing the homogeneity condition ($\sum_j \gamma_{ij} = 0$) on equation (9) and by using an F test to compare the residual sum of squares before and after its imposition.

Comparison of the LES and AIDS

Before reporting the empirical results, we briefly discuss some of the properties of the elasticities of the two demand systems. The expenditure and uncompensated own-price elasticities for the LES are:

$$\eta_i = \theta_i / w_i \qquad (10)$$

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[6] The term α_o can be interpreted as the outlay required for a minimal standard of living when prices are equal to 1 as in a base year (*10*, p. 316).

[7] One of the adding up restrictions is redundant when the homogeneity and Slutsky symmetry conditions are imposed. That is, if $\sum_i \gamma_{ij} = 0$ and $\gamma_{ij} = \gamma_{ji}$, then $\sum_i \gamma_{ij} = \sum_i \gamma_{ji} = \sum_j \gamma_{ij} = 0$.

[8] For example, for 4 commodities, there are 12 unknown parameters; for 8 commodities, there are 42 unknown parameters; and for 12 commodities, there are 88 unknown parameters.

[9] There is an econometric problem in the linear approximate version. If $\log P^*$ is not treated exogenously, the dependent variable, w_i, appears on both sides of the equation and the resultant estimators will not necessarily possess desirable sampling properties. However, following Deaton and Muellbauer (*10*), we ignore this econometric problem in obtaining parameter estimates.

and:

$$\epsilon_{ii} = -1 + (1 - \theta_i)p_i\mu_i/w_iY \qquad (11)$$

respectively. For the AIDS, the expenditure and uncompensated own-price elasticities are given by:

$$\eta_i = 1 + \beta_i/w_i \qquad (12)$$

and:

$$\epsilon_{ii} = -1 + [\gamma_{ii} - \beta_i(\alpha_i + \sum_k \gamma_{ik} \log p_k)]/w_i \qquad (13)$$

respectively. With regard to changes in the expenditure elasticities corresponding to changes in the ith budget share, the LES reflects the property that expenditure elasticities become more elastic as the ith budget share decreases; that is, $\partial\eta_i/w_i^2 - \theta_i/w_i^2 < 0$, as marginal budget shares are always restricted to be positive. The implication is that as the budget share for a necessary commodity, such as food, decreases (which it has over time), its expenditure elasticity increases (assuming no inferior goods). This hypothesis seems unrealistic. However, the AIDS and the LA/AIDS—as neither restricts marginal budget shares to be positively valued—allow the expenditure elasticity to decrease with respect to a decrease in the budget shares for necessities ($\beta_i < 0$). Mathematically, $\partial\eta/\partial w_i = -\beta_i/w_i^2 > 0$ for $\beta_i < 0$. Thus, in this situation, the AIDS and LA/AIDS possess a more desirable property than the LES. Concerning the properties of the own-price elasticities with respect to a change in w_i, in the LES, $\partial\epsilon_{ii}/\partial w_i = -(1 - \theta_i)p_i\mu_i/w_i^2Y < 0$, assuming $0 < \theta_i < 1$ and $\mu_i > 0$. Thus, as the ith budget share decreases, the own-price elasticity becomes more inelastic, as expected. In the AIDS, the sign of $\partial\epsilon_{ii}/\partial w_i$ depends on the relative magnitudes of γ_{ii} and $\beta_i(\alpha_i + \Sigma\gamma_{ik} \log p_k)$ (see equation (13)). *A priori,* it is extremely difficult to assign a positive or negative value to the change in ϵ_{ii} with respect to a change in the budget share, w_i.

Estimation of Models

To estimate the demand models, one must add an error term, e_{it}, to each equation. The stochastic specification for the disturbance terms is assumed to have zero mathematical expectation, to be temporarily uncorrelated, and to have a contemporaneous variance-covariance matrix Ω. Problems arise in both the LES and the AIDS because the sum of the budget shares equals 1. In this case, the variance-covariance matrix is singular. If no autocorrelation is present, one can apply FIML procedures by arbitrarily deleting an equation (see (*1*, *4*)).

We used the TSP program by Hall and Hall (*13*) and discussed in Berndt, and others (*3*) to obtain FIML estimators of the parameters for both the LES and the AIDS and OLS estimates for the LA/AIDS. The term α_0 was assigned *a priori* to be the cost at base year prices. This value was equal to $586.90 in the base year 1972. Also, following Deaton and Muellbauer (*10*, p. 316), log P was approximated by Stone's index $\log P^* = \sum_k w_k \log p_k$. As already discussed, the use of this approximation simplifies the estimation procedure considerably; however, not without some cost.

Data

We used annual U.S. time series data for 1948-78 to estimate the three models. For the four food groups, the commodities are the following: meats (beef and veal, pork, fish, and poultry), fruits and vegetables, cereal and bakery products, and miscellaneous foods (dairy products, eggs, imported sugar, and some minor items). Manser (*18*) used similar commodity classifications.

The primary source for these data is the U.S. Department of Agriculture (USDA) series called consumer expenditures on domestic farm food products bought by civilians (*23*, *24*, *25*). The USDA series is available for seven commodity groups and excludes fish and imported foods. To obtain our meats group, we adjusted the USDA meat series which includes beef, veal, and pork to include fish and poultry by reconstructing these expenditures according to the method used by Christensen and Manser (*8*). Data for fruits and vegetables were taken directly from the USDA series. Grain mill and bakery products were aggregated into the cereal and bakery products group. Imported foods were a negligible component of both the fruits and vegetables and cereal and bakery products groups. However, imports of sugar were found to be significant. Imports of sugar along with the expenditure series for dairy products, (constructed) eggs, and USDA's other food products were combined into a catch-all miscellaneous foods group. The price series are the published consumer price indexes for meat, poultry, and fish; fruits and vegetables; and cereal and bakery products. We created an implicit price deflator for the miscellaneous foods group by dividing current dollar expenditures by their constant (1972) dollar counterpart. (See (*7*) for a more detailed listing of these data sources.)

Empirical Results

For the four food groups, we used FIML techniques to obtain estimates of the parameters of the LES and the AIDS, whereas we used the OLS technique to estimate the linear approximate AIDS using Stone's index. Table 1 gives the results of the LES with food expenditure and own-price estimated elasticities reported in columns four to eight.[10] The estimated food expenditure elasticities for the LES model indicate that two of the four commodities are relative luxuries; that is, food expenditure elasticities are greater than 1 for meats and miscellaneous foods. Fruits and vegetables are relatively inferior, and cereal and bakery products are relative necessities. The estimated own-price elasticities indicate relatively inelastic demand for all groups, except fruits and vegetables. Referring to the Pigou relation, we observe the proportionality variable, ϕ, is approximately 0.7, implying that the estimated own-price elasticity is about 70 percent of the estimated expenditure elasticity. We obtained these values by using the approximation formula, $\epsilon_{ii} \approx \phi\eta_i$.

Table 2 reports estimates for the AIDS. First, note that the estimated expenditure elasticities differ greatly between the AIDS and the LES. Here, meats and fruits and vegetables are relative luxuries, and cereal and bakery products and miscellaneous foods are relative necessities. All the estimated own-price elasticities indicate relatively inelastic demand. Calculation of the Pigou relation for this system reveals that no approximate proportional relationship exists between price and expenditure elasticities. These estimates appear more reasonable than their LES counterparts.

Finally, table 3 contains results for the approximate version of the AIDS, with and without homogeneity imposed. The magnitude of most of the intercept and expenditure coefficients is substantially higher. This holds for the associated t-values as well. F-values indicate that homogeneity is rejected for meats and miscellaneous foods.

A comparison of the homogeneous nonsymmetric approximate (table 3, all columns with H boxheads) model results with the full AIDS system results

Table 1—Linear expenditure system (LES): Estimates for four food groups

Food group i		Estimated coefficients[1]		Elasticities[2]						Pigou relationship ϕ
		Marginal budget share θ_i	Minimum subsistence level μ_i	Expenditure		Uncompensated price				
				Total[3]	Food	Meats	Fruits and vegetables	Cereal and bakery products	Miscellaneous foods	
Meats	(1)	0.537 [4](12.5)	−18.738 (−.7)	0.756	1.738	−1.049	−0.426	−0.100	−0.137	0.604
Fruits and vegetables	(2)	−.078 (−1.0)	143.443 (5.2)	−.169	−.389	−.013	.319	.022	.031	.820
Cereal and bakery products	(3)	.117 (5.3)	31.602 (4.7)	.379	.871	.029	−.219	−.614	−.071	.705
Miscellaneous foods	(4)	.424 (3.7)	44.063 (3.0)	.520	1.196	.040	−.304	−.071	−.867	.725

[1] Coefficients are based on U.S. data from the years 1948 to 1978.
[2] Elasticity formulas are calculated at mean (1948-78) values.
[3] Based on first-stage expenditure elasticity for food of 0.435.
[4] Values in parentheses are asymptotic t-statistics.

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[10] The total income elasticities in table 1, column 4, are obtained by a simple conversion formula (see (*5*, p. 26)). Similar (yet not so simple) concepts hold for price elasticities. For meat, for example, the estimated income elasticity with respect to total expenditures or income is (0.435) (1.738) = 0.756, where 0.435 is the estimated income elasticity for food with respect to all expenditures from the first stage (by use of a LES system) and 1.738 is the estimated expenditure elasticity for meats with respect to total expenditures for food.

Table 2—Almost ideal demand system (AIDS): Estimates for four food groups

Food group i		Estimated coefficients[1]					
		α_i	β_i	γ_{i1}	γ_{i2}	γ_{i3}	γ_{i4}
Meats	(1)	0.327	0.328	0.110	−0.140	−0.012	0.042
		[2](111.8)	(8.7)	(4.6)	(−9.3)	(−1.2)	(1.1)
Fruits and vegetables	(2)	.209	.052	−.140	.160	−.004	−.016
		(88.2)	(1.3)	(−9.3)	(4.4)	(−.2)	(−.3)
Cereal and bakery products	(3)	.129	−.078	−.012	−.004	.017	−.001
		(82.3)	(−4.1)	(−1.2)	(−.2)	(1.0)	(−.0)
Miscellaneous foods	(4)	.336	−.302	.042	−.016	−.001	−.026
		(72.9)	(−4.5)	(1.1)	(−.3)	(−.2)	[3](−.3)

Food group i		Elasticities[4]						Budget share, 1948-78 average value	Pigou relationship ϕ
		Expenditure		Uncompensated price					
		Total[5]	Food	Meats	Fruits and vegetables	Cereal and bakery products	Miscellaneous foods		
Meats	(1)	0.897	2.062	−0.992	−0.672	−0.178	−0.220	0.309	0.481
Fruits and vegetables	(2)	.055	1.260	−.780	−.263	−.053	−.170	.202	.209
Cereal and bakery products	(3)	.183	.421	.098	.090	−.800	.190	.134	1.900
Miscellaneous foods	(4)	.064	.147	.399	.131	.109	−.787	.355	5.354

[1] Coefficients are based on U.S. data for 1948-78.
[2] Values in parentheses are asymptotic t-statistics.
[3] This is an approximate t-value as there are no covariance terms.
[4] Elasticity formulas are calculated at mean (1948-78) values.
[5] Based on first-stage expenditure elasticity for food of 0.435.

(table 2) reveals little diversity in the expenditure coefficients (β_i) and in some price coefficients, such as γ_{i1} and γ_{i4}, but large differences in the intercepts, α_i, and in the γ_{i2} and γ_{i3} estimates. Because of the similarity in the β_i's, the food expenditure elasticity results are approximated exceedingly well by the linear version. The own-price elasticities do not indicate such a high degree of similarity. However, all but the fruits and vegetables own-price elasticities in the approximate version are nearly the same value as in the complete AIDS.[11] Again, the Pigou relation is not evident in the LA/AIDS.

Comparison of the results of either of these two models with the results of the LES indicates even greater differences. First, one should note that the proportional relationship between the expenditure and own-price elasticities holds for all groups of the LES. The AIDS does not possess this proportionality relationship and shows higher expenditure elasticities for all groups except cereal and bakery products and miscellaneous foods and shows lower own-price elasticities for all groups except fruits and vegetables and cereal and bakery products.

Conclusions

This analysis demonstrates that the AIDS of Deaton and Muellbauer (*10*, *11*) is a viable system for analyzing the demand for food commodities. The AIDS avoids the unrealistic approximate proportionality relationship between income and own-price elasticities that the LES may exhibit. The AIDS also has some advantages over the LES in that income elasticities can decrease as budget shares decrease for necessities such as food.

As a first-order approximation to a complete demand system, the linear approximate version with homogeneity imposed performs reasonably well with respect to estimated magnitudes of elasticities. The advantage of the approximate version is its ease of estimation; theoretically, however, no claims can be made with respect to the properties of its estimators.

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[11] The Stone index is a good approximation of log P.

Table 3—Effects of relaxing the homogeneity condition in the static linear approximate almost ideal demand system ((LA/AID) (nonsymmetric): Estimates for four food groups

Food group i		Estimated coefficients[1]									
		α_i^*		β_i		γ_{i1}		γ_{i2}		γ_{i3}	
		H[2]	NH[3]	H	NH	H	NH	H	NH	H	NH
Meats	(1)	−1.763 [4](−6.1)	−0.564 (−1.5)	0.328 (7.2)	0.140 (2.4)	0.106 (4.7)	0.120 (6.6)	−0.118 (−2.3)	−0.042 (−1.0)	−0.048 (−1.1)	−0.056 (−1.6)
Fruits and vegetables	(2)	−.191 (−.9)	.230 (.7)	.062 (1.8)	−.004 (−.1)	−.131 (−7.6)	−.127 (−7.4)	.126 (3.3)	.153 (3.7)	.030 (.9)	.028 (.9)
Cereals and bakery products	(3)	.553 (4.3)	.538 (2.5)	−.067 (−3.3)	−.064 (−1.9)	−.005 (−.5)	−.005 (−.5)	−.025 (−1.1)	−.026 (−1.0)	−.031 (1.6)	.032 (1.6)
Miscellaneous foods	(4)	2.424 (6.8)	.787 (1.9)	−.328 (−5.8)	−.070 (−1.1)	.030 (1.0)	.011 (.7)	.020 (.3)	−.084 (−1.7)	−.013 (.2)	−.003 (−.1)

Food group i		γ_{i4}		$\sum_j \gamma_{ij}$		Elasticities[5]					
						Expenditure		Uncompensated price			
						Food		Meat		Fruits and vegetables	
		H	NH	H	NH	H	NH	H	NH	H	NH
Meats	(1)	0.060 (.6)	0.010 (0.4)	0 0	0.032 (4.2)	2.062	1.452	−1.006	−0.758	−0.594	−0.227
Fruits and vegetables	(2)	−.025 (−.4)	−.043 (−1.7)	0 0	.011 (1.6)	1.310	.982	−.753	−.622	−.486	−.239
Cereals and bakery products	(3)	−.001 (−.1)	−.001 (−.1)	0 0	0 (.1)	.515	.522	.125	.114	−.084	−.094
Miscellaneous foods	(4)	−.037 (−.3)	.032 (1.0)	0 0	−.043 (−5.0)	.087	.802	.388	.096	.388	.197

Food group i		Elasticities				Pigou relationship ϕ		Homogeneity test results
		Uncompensated price						F-value[6] critical value F(1,25) = 4.24
		Cereal and bakery products		Miscellaneous foods				
		H	NH	H	NH	H	NH	
Meats	(1)	−0.295	−0.239	−0.162	−0.124	0.488	0.522	17.40*
Fruits and vegetables	(2)	.111	.140	−.229	−.206	.332	.243	2.43
Cereal and bakery products	(3)	−.702	−.704	−.154	.157	1.390	1.349	.01
Miscellaneous foods	(4)	.084	−.018	−.794	−.843	1.031	1.051	14.78*

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[1] Coefficients are based on U.S. data for 1948-78.

[2] H indicates results from the homogeneous model.

[3] NH indicates results from the nonhomogeneous model.

[4] Coefficients in parentheses are t values.

[5] Elasticities are calculated at mean (1948-78) values.

[6] * indicates rejection of the homogeneity hypothesis.

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Summary Statistics and Forecasting Performance

By Don Larson*

Abstract

The simultaneous equation econometric model has recently come under increasing attack as a policy analysis and forecasting tool. However, traditional methods of choosing among competing models rely heavily on the use of summary statistics. It is shown, by example, that choosing a model with relatively better summary statistics does not guarantee getting the best unconditional out-of-sample forecasts. Other methods of forecasting time series that allow relative evaluation of the out-of-sample forecasting ability of econometric simultaneous equation models are also examined.

Keywords

Summary statistics, econometric models, autoregression, ARIMA, identification

Introduction

The simultaneous equation econometric model has recently come under increasing attack as a policy analysis and forecasting tool. The criticisms are based on both theoretical results and the more pragmatic findings of *ex ante* forecast comparisons. The criticisms lead to questions concerning not only the efficiency of out-of-sample forecasts but also the developmental methodology of most standard multivariate models.

Despite the accumulating doubts surrounding such models, a variety of firms exist and prosper by selling the results of their econometric models—a fact which attests to the perceived usefulness of their models. While there are a variety of uses to which econometric models might be put other than forecasting, the ability to replicate history is often the final measure of a model's validity.

The presence of a large commercial market in econometric models and forecasts is consistent with a growing list of practical and theoretical modeling difficulties. In fact, such problems may help explain the specialization of forecasting firms.

Given the combination of doubt and usefulness associated with econometric models, the evaluation of competing models and methodologies becomes doubly important. Persons in both the public and private sectors are required, with increasing frequency, to evaluate econometric models or economic analysis based on models or modeling techniques.

The traditional method of quickly evaluating the accuracy of such models relies heavily on the use of summary statistics, such as t-scores, F-tests, R^2's, and Durbin-Watson (D-W) statistics, and these statistics represent the type of information generally requested and provided by the major economic journals. One purpose of this article is to discuss the limited information provided by such statistics and to offer an illustration of those limits. Another related purpose is to discuss some easily available, alternative methods which, through comparison, can provide evaluations of a model's relative forecasting abilities.

I will review earlier criticisms of fixed parameter econometric models, present an *ad hoc* quarterly model of the corn sector along with estimation results from the model, discuss possible misinformation provided by initial model results and provide an alternative estimation, discuss several alternative methods of unconditional forecasting, and, finally, compare out-of-sample results.

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Model Limits

There is an abundance of economic literature that discusses the gaps between modern economic theory and the structure of a simultaneous equation econometric model estimated from time series data. Such problems supersede the practical difficulties of data definition and collection, along with estimation efficiency, and relate to the very logic by which economic behavior is expressed in terms of a fixed-parameter mathematical model. Sims (*30*), in particular, provides an excellent review of such problems.[1]

Perhaps the best known dilemma facing modelers who use time series data centers on Lucas' (*20*) argument that policy variables normally considered exogenous and independent actually determine the way in which each of the variables relate to all others. His conclusion is based on Muth's (*23*) assertion that "[t]he way expectations are formed depends specifically on the structure of the relevant system describing the economy." Lucas argues that as policy changes are effectively incorporated into expectations, new policy results in new rules of response so that the reduced forms of econometric models do not have fixed parameters. Prescott and Kydlund (*26*) have shown, in turn, that policy formation without regard to expectations formation can lead to what Sims has termed "Peter White" policymaking,[2] which, by necessity, must always go awry.

Lucas' results lead to the dilemma that no economically or politically determined variable can truly be considered exogenous. More recently, "causality" tests have been used to identify exogenous/endogenous relationships, but on a practical level, the tests themselves must be performed on a reduced form consistent with a multitude of theoretical structures. See, for example, Sims (*28, 29, 31*), Granger and Newbold (*15*), Bishop (*4*), Ashley, Granger and Schmalensee (*1*), Barnett, Bessler, and Thompson (*2*).

Should the problem of exogeneity be solved, Sims (*30*) has suggested that yet another identification problem exists. Sims cites the work of Hatanaka (*16*), who reevaluated Sargan's (*27*) conclusions concerning the simultaneous equation identification problem in models containing lagged dependent variables and autocorrelated residuals. Hatanaka found that if the lag length and lag distribution are not known *a priori*, but rather are determined in the estimation process, then the identification rule is changed. In order to identify an equation, at least one strictly exogenous variable must be located in each of the other equations of the model. Repeat occurrences of the same variable with different lags in a single equation do not qualify as "strictly exogenous."

The application of Hatanaka'a criteria leaves many models underidentified. Sims also notes that the exogeneity of variables is often determined by convenience. While this does limit the size of the model, it further exacerbates the confusion between exogenous/endogenous variables and may limit the number of variables actually available to identify the system.

Given these problems, it is not surprising that simultaneous equation models have had limited success as forecasting tools.[3] Gordon (*12*) has drawn on the work of McNees (*22*) to document the failure of macroeconomic forecasters in the seventies. More recently, Just and Rausser (*17*) have presented evidence that commodity price forecasts are generally better provided by the futures market than by large-scale econometric models. Their findings are consistent with a growing list of works describing the consistency between futures markets and Muth's hypothesis (see (*8*), (*10*)).

As already mentioned, model forecasts do exist and prosper in the market place, and it is this marketability which, in fact, documents their usefulness. At the same time, there are a variety of models from which to choose and a number of potential shortcomings to avoid. In the following section, I will argue that it is generally difficult to evaluate a model from the usual summary statistics provided by vendors and by most economic journals. To illustrate this point, I provide and evaluate a quarterly corn model.

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[1] Italicized numbers in parentheses refer to items in the References at the end of the article.

[2] Peter White will ne'er go right/Would you know the reason why?/He follows his nose where'er he goes/And that stands all awry.—Nursery Rhyme.

[3] Criticisms of particular models over particular periods are by no means a recent phenomenon (see Christ (*7*)); however, as marketed forecasts have come of age, historical evaluations of their accuracy have become available.

The Model

Consider the following quarterly model:

$$\begin{aligned}
CPLANT_t &= f_1(CAL_t, CP/SMP_{t-3}, \\
&\qquad CP/SMP_{t-4}, SGAL_t) && \text{(1a)} \\
CHARV_t &= f_2(CPLANT_t) && \text{(1b)} \\
CDEMD_t &= f_3(BEEFPR_t, CHKPR_t, CP/IN_t, \\
&\qquad SMP/IN_t, WHP/IN_t, D2, D3, D4) && \text{(1c)} \\
CSTKS_t &= CSTKS_{t-1} + CPR_t + CIMP_t \\
&\qquad - CEXP_t - CDEMD_t && \text{(1d)} \\
CPR_t &= CHARV_t * CYLD_t && \text{(1e)}
\end{aligned}$$

where variable definitions are provided in table 1.

The model ignores several important structural aspects of the corn market, and prices are exogenous. However, the model is similar to some large commercial agricultural forecasting models, and it serves well in illustrating several points about summary statistics.

Equation (1a), (1b), and (1e) are annual equations when the model is simulated as CPLANT and CHARV and CPR are nonzero only one quarter of the year. As a result, seasonal dummies have been included only in equation (1c). Production and stocks are solved by identities. When performing forecasts greater than three quarters, one often takes yields for the type of model above from trend levels, so that yields become a function of time or lagged yields. However, as the time horizon shortens, the U.S. Department of Agriculture's (USDA) forecasted yields, based on weather and/or survey data, are often used.

A linear stochastic form of the three behavioral equations was estimated by use of a three-stage least squares procedure on quarterly data from USDA's T-DAM data base for the 1960-76 period. Data for the 1977-79 period were retained for out-of-sample tests. Table 2 shows the first stage results. Third-stage results are available from the author upon request.[4]

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[4] From a theoretical point of view, there are several arguments for choosing one estimation process over another. The production and sale of corn is a recursive process, a fact which is often evoked when one justifies a set of OLS estimates as unbiased. However, reported data are often reviewed and adjusted by a USDA committee to provide consistency across aggregates. Although the adjustments represent a logical necessity for data presentation, the changes potentially disrupt predetermination among the series, justifying a three-stage least squares, or seemingly unrelated estimation procedure. Both estimation techniques were used, and they resulted in only minor differences.

Table 1—Variable definitions

Variable	Definition	Unit
Endogenous:		
CPLANT	Planted acreage of corn	1,000 acres
CHARV	Harvested acreage of corn	do.
CDEMD	Domestic disappearance of corn	Million bushels
CSTKS	Corn stocks	do.
CPR	Corn production	do.
Exogenous:		
CAL	Acres allocated to corn	1,000 acres
SGAL	Acres allocated to sorghum	do.
CP/SMP	Ratio of corn farm price to soy meal price	Dollars/bushel
BEEFPR	Beef production	Million pounds
CHKPR	Broiler production	do.
CP/IN	Ratio of corn farm price to real disposable income	Dollars/bushel over billion 1972 dollars
SMP/IN	Ratio of soy meal price to real disposable income	do.
WHP/IN	Ratio of wheat farm price to real disposable income	do.
CIMP	Corn imports	Million bushels
CEXP	Corn exports	do.
CYLD	Corn yield	Bushels/acre

Table 2—First-stage results: Original series

Equation

$$\text{EQ1: CPLANT} = \underset{(6.47)}{40608.12} + \underset{(4.93)}{0.054*\text{CAL}} + \underset{(3.20)}{828929*\text{CP/SMP3}} - \underset{(-0.96)}{298507*\text{CP/SMP4}} + \underset{(0.32)}{1.05*\text{SGAL}}$$

D-W D Statistic = 2.0

F Ratio	6782.80
Prob>F	0.0001
R-Square	0.9984

$$\text{EQ2: CHARV} = \underset{(0.74)}{978.83} + \underset{(44.1)}{0.84*\text{CPLANT}}$$

D-W D Statistic = 2.0

F Ratio	296368.19
Prob>F	0.0001
R-Square	0.9999

$$\text{EQ3: CDEMD} = \underset{(3.07)}{415.48} + \underset{(1.00)}{0.067*\text{BEEFPR}} + \underset{(2.02)}{0.28*\text{CHKPR}} - \underset{(-4.26)}{1230345*\text{CP/IN}} + \underset{(0.93)}{1422.1*\text{SMP/IN}} + \underset{(3.83)}{561138*\text{WHP/IN}} - \underset{(-10.9)}{509.8*\text{D2}} - \underset{(-0.94)}{43.0*\text{D3}} + \underset{(3.30)}{107.7*\text{D4}}$$

D-W D Statistic = 2.07

F Ratio	61.48
Prob>F	0.0001
R-Square	0.9044

Note: t-ratios are given in parentheses.

For the first-stage statistics, the distinctive features of the equations as a group are that they possess remarkably high R^2s and exceedingly solid Durbin-Watson (D-W) statistics.[5] All the parameter signs match *a priori* expectations, and most of the t-scores are significant at standard levels. Third-stage results differ only slightly from the first-stage results.

Examining the residuals in the equation for domestic disappearance, however, reveals an autocorrelation problem (table 3). Although the sample correlation coefficient at lag 1 confirms the earlier D-W statistic, problems arise at lags 2, 3, and particularly 4. Given the seasonal nature of agriculture and the fact that quarterly data are used, it is not surprising to find signs of fourth-order serial correlation. Although the potential problem is well known and Wallis (*32*) has devised a measure similar to the D-W statistic to detect its presence, such information is seldom provided to purchasers or users of existing models.

The Effects of Autocorrelation

It is well known that regression coefficients are less efficient, but unbiased, when estimated in the presence of an autocorrelated error structure (*18*, p. 275). Moreover, the multiple correlation coefficient increases in the presence of autoregressive residuals. As Bishop (*5*, p. 14) notes, Granger and Newbold have been particularly critical of reporting high R^2's under such circumstances:

> In time series regressions involving the levels of economic variables, one frequently sees coefficients of multiple correlation (R^2) much higher than 0.9. If these indicate anything at all, they presumably imply an extremely strong relationship between the dependent variable and the independent variables. This is extremely misleading on many occasions, as comments noting poor forecast performance which sometimes follow these equations will testify. In fact, the high R^2

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[5] Given the structure of the model, variables which have a value only once every four quarters will, of course, always have a D-W value of 2, given the formula:

$$d = \frac{\Sigma\,(e_t - e_{t-1})^2}{\Sigma\,(e_t)^2}$$

Table 3—Sample autocorrelation coefficients on CDEMD residuals: Original series

Lag	Correlation	−1 9 8 7 6 5 4 3 2 1	0 1 2 3 4 5 6 7 8 9 1
0	1.00000		*******************************
1	−.04088	. *	.
2	−.23487	*****	.
3	−.22213	.******	.
4	.34186	.	**********
5	−.14682	. ***	.
6	−.07209	. *	.
7	−.08108	. **	.
8	.10161	.	** .
9	−.00398	.	.
10	−.13777	. ***	.
11	−.13508	. ***	.
12	.05048	.	* .
13	.15769	.	*** .
14	−.01194	.	.
15	−.08283	. **	.
16	−.01853	.	.
17	.21537	.	**** .
18	.02823	.	* .
19	−.18358	. ****	.
20	−.04120	. *	.

Note: *marks correlations; • marks standard error.

values could be no more than a reflection of the fact that the dependent variable is highly autocorrelated and could easily be achieved simply by regressing the variable on its own past. Thus, in such circumstances, the value of R^2 says nothing at all about the strength of the relationship between the dependent and independent variable.

As the multiple correlation coefficient has proven an unreliable tool in applied econometrics, emphasis has been placed on other tests of significance, primarily t-scores. Bishop (*5*, p. 13) notes, however:

> What is also well documented in the literature, but often overlooked in practice, is that the usual tests of significance, when performed in the presence of autocorrelated errors, are biased. For example, if positive first-order autocorrelation is present in the error structure and the independent variable is also autocorrelated, the estimates of the standard errors on each of the coefficients (s_b) will be biased downward in most situations. When the standard error of the coefficient is underestimated, the t-statistic on that coefficient is obviously overstated as it is computed as $t = (b-b_o)/s_b$, implying greater explanatory power for that variable than actually exists. This situation can easily lead to the inclusion of a statistically irrelevant variable in the final model. If the error structure exhibits negative serial correlation and the independent variable is positively autocorrelated, the standard errors of the coefficients are likely to be overestimated, possibly leading to the elimination of a statistically significant variable from the model.

The result of Bishop's caveat is that, unless additional information concerning residuals is provided, the relevance of any set of t-scores remains unknown. Let the buyer beware.

Fortunately for the modeler, simple differencing usually greatly reduces the serial correlation problem. To illustrate, the following transformation was made of the model's original data:

$$x^*_t = x_t - x_{t-4} \qquad (2)$$

so that quarterly observations now represent year-to-year changes between quarters.

With the exception of intercept values and the error terms, the transformation does not structurally affect the expected values of the estimated coefficients. However, by reducing the serial correlation problem, the estimation process should prove more efficient. Table 5 shows the first-stage results.

In terms of first-order autocorrelation, the transformation appears a failure. The D-W statistic, which is the statistic most often provided relating to serial correlation, dropped from 2.07 to 1.66 in the disappearance equation, which is still indeterminate at a 5-percent level of confidence. Table 4, however, provides the sample autocorrelation coefficients of the residuals which, in fact, show no sign of serial correlation. In addition, I performed a collective test on the errors by utilizing the sample autocorrelation coefficients across 24 lags to provide the Box-Pierce Q-statistic, which has an approximate Chi-square distribution (see Nelson (*24,* p. 115)). As a result of the transformation, the Q-statistic dropped from 28.04 to 18.53. The confidence with which the series could be termed "white noise" increased from 21 percent to 76 percent.

Transformation of the data requires a transformation in the form of the model. A final measure of the appropriateness of the transformation is to test whether the seasonal parameters dropped are statistically different from zero when retained and estimated with transformed data.

An F-test was performed to test the hypothesis that the coefficients on the seasonal dummy variables collectively equal zero. The hypothesis could not be rejected for any of the equations. F-score results are reported in table 6.

The reduction in serial correlation had few meaningful effects on the summary statistics. As already mentioned, the D-W statistic on the demand equation dropped slightly into the indeterminate range. The changes on the t-scores were mixed. Some scores, such as the score on the soy meal price variable in the demand equation, improved, whereas others, such as the scores on the beef and chicken production coefficient, declined.

The multiple correlation coefficients all dropped dramatically; however, the two sets of numbers are

Table 4—Sample autocorrelation coefficients on CDEMD residuals: Transformed series

```
Lag Correlation   -1  9   8   7   6   5   4   3   2   1  |  0   1   2   3   4   5   6   7   8   9   1
  0      1.0000                                          |  ***********************************
  1       .1607                                 .        |  ***       .
  2      -.1535                                    ***   |            .
  3      -.1488                               . .   ***  |              .
  4      -.1833                               .    ****  |              .
  5      -.0763                               .      **  |              .
  6       .2305                               .          |  *****       .
  7       .2055                               .          |  ****        .
  8      -.1445                               .     ***  |              .
  9      -.0525                               .       *  |              .
 10      -.1034                               .      **  |              .
 11      -.0070                               .          |              .
 12      -.0032                               .          |              .
 13       .0582                           .              |  *             .
 14       .0302                           .              |  *             .
 15      -.0007                           .              |                .
 16      -.0313                           .           *  |                .
 17       .1532                           .              |  ***           .
 18       .0697                           .              |  *             .
 19      -.1503                           .         ***  |                .
 20      -.0360                           .           *  |                .
```

Note: *marks correlation; . marks standard error.

Table 5—First-stage results: Transformed series

Equation

$$\text{EQ1: CPLANT} = \underset{(2.59)}{0.036{*}\text{CAL}} + \underset{(4.54)}{705924{*}\text{CP/SMP3}} - \underset{(-2.65)}{191553{*}\text{CP/SMP4}} + \underset{(9.27)}{0.17{*}\text{SGAL}}$$

D-W D Statistic = 2.0

F Ratio	32.44
Prob>F	0.0001
R-Square	0.7100

$$\text{EQ2: CHARV} = \underset{(28.4)}{0.87{*}\text{CPLANT}}$$

D-W D Statistic = 2.0

F Ratio	809.39
Approx PR>F	0.0001
R-Square	0.9353

$$\text{EQ3: CDEMD} = \underset{(0.25)}{0.016{*}\text{BEEFPR}} + \underset{(1.97)}{0.30{*}\text{CHKPR}} - \underset{(-3.43)}{770789{*}\text{CP/IN}} +$$
$$\underset{(1.40)}{1592.1{*}\text{SMP/IN}} + \underset{(3.22)}{396995{*}\text{WHP/IN}}$$

D-W D Statistic = 1.65

F Ratio	6.81
Prob>F	0.0001
R-Square	0.3956

Note: t-ratio given in parentheses.

Table 6—F-scores on hypothesis that coefficients on seasonal dummies collectively equal zero

Equation	F-score	Value at 99-percent confidence level
Acres planted	0.04	7.1
Acres harvested	.58	7.2
Total disappearance	.13	4.2

not comparable, and they reveal very little about the relative efficiency of either model.

Finally, the F-scores associated with each of the equations fell substantially with the data transformation. However, the hypothesis that the equations contain no explanatory power could still be rejected overwhelmingly.

Forecasting

Although a reduction in the level of serial correlation should improve the parameter estimates for any particular model, questions remain unanswered concerning the reliability and efficiency of the model's forecasts. In this section, I describe three alternative methods of forecasting time series, each of which requires less information, and can thus be more quickly estimated and forecasted than the model in the second section. In the last section, I compare out-of-sample forecast performance for these models and for the restricted equation model.

ARIMA Model

A general class of auto regressive-integrated-moving-average (ARIMA) models can supply quick and often relatively efficient forecasts for many time series (*6*). For the most part, ARIMA models are purely statistical and impart little information about the economic processes involved. Future innovations in a series are modeled on past innovations, allowing a univariate model form and a self-contained forecast of the time series.

The general form of the model, ARIMA(p,d,q) is:

$$w_t = b_1 w_{t-1} + \ldots + b_p w_{t-p} + u_t - c_1 u_{t-1} - \ldots - c_q u_{t-q} \quad (3)$$

where w is a form of the original series differenced d times, u is an error term, and b and c are fixed coefficients.

Box and Jenkins provide a three-stage modeling process of: (1) identifying potential models, (2) estimating one or more models, and (3) testing the restriction each form places on the model's parameters. With the advent of the appropriate computer software, ARIMA models can be quickly and cheaply estimated and forecasted.

Vector Regression

Sims (*30,* p. 14) has noted:

> Much recent theoretical work gives rigorous foundation for a rule of thumb that in high dimensional models restricted parameters can easily produce smaller forecast or projection errors than unrestricted estimators, even when restrictions are false . . . . Thus models whose self-proclaimed behavioral interpretation is widely disbelieved may nonetheless find satisfied users as tools of forecasting and policy projection.

As a possible solution, Sims proposed the estimation of a reduced version of the model, treating all variables as endogenous, without imposing any *a priori* restrictions on the parameters. Restrictions could then be added and tested in a more systematic manner.[6] Sims termed this process vector autoregression.

Although models tend to become large quickly under Sims' methodology, vector autoregression does provide a readily available standard which can be applied to more restricted model forms. The functional form of a vector regression system is simply:

$$X_t = P(L)X_{t-1} + Z_t \quad (4)$$

where X is a vector of endogenous variables, L represents a lag operator whose length is determined by the data, P is a matrix of estimated parameters, and Z is a vector of disturbance terms.

Although the choice of variables may be guided by economic theory, the general form of the model is most likely consistent with a variety of competing theoretical models, and as such, trades the potential efficiencies of a parsimonious parameterization for the ability to test particular restrictions on a more general form.

Autoregressive Open Model

A third alternative to the standard simultaneous equation system has been recently used by Lamm (*19*) for the U.S. food and agricultural sector. The method is similar to Sims' method with the exception that the distinction between exogenous and endogenous variables is retained.

This distinction helps reduce the size of the model and links the modeled sector to relevant information outside the sector.

The model is written as:

$$X_t = P(L)X_{t-1} + H(L)Y_t + U_t \quad (5)$$

where X is again a vector of endogenous variables, Y is a vector of exogenous variables, U is a vector of disturbance terms, L is the lag operator, and P, H are vectors of estimated parameters.

The form of the model allows for checks on any set of restrictions across parameters as well as for checks on the assumption of exogeneity.

Out-of-Sample Forecasts

In this section, I present the out-of-sample forecast performance for five methodologies:

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[6] As Malinvaud (*21*) notes, the assumptions concerning efficiency gains implicit in restricted-form equation systems, with few exceptions, have gone untested in most models since the pioneering work of Frisch and Tinbergen.

1. The restricted simultaneous equation system presented in the second section, estimated with the original time series;
2. The same model estimated with seasonally differenced data as described in Section 3;
3. ARIMA models for each series;
4. A vector autoregression model; and
5. An autoregressive open model.

The forecasts were performed as follows:

1. Each model was estimated from a sample, and then resulting parameters were used to forecast one, two, three, and four periods ahead.
2. One period was then added to the sample, and the process in the first step was repeated until the observation set had been exhausted.

Because the emphasis here is on a quarterly model, only the variables for which quarterly values exist, CSTKS and CDEMD, were forecast. In order not to penalize the more fully specified models, I used actual values for all endogenous variables other than CSTKS and CDEMD when forecasting these two variables.

Table 7 shows the mean absolute error (MAE) associated with each methodology for each of the forecast periods when CDEMD is forecast. Table 8 provides the same information for forecasts of CSTKS.

The most dramatic result of the out-of-sample simulation is the more than threefold reduction in the MAE in the restricted equation models due to simply differencing the data series. This reduction occurred despite the fact that the differenced model had a lower R^2, a mixed set of t-scores, and a less impressive D-W statistic. In short, the summary statistics had provided all the wrong signals.

Once the differencing occurred, the restrictions did indeed seem to improve the model's forecasting efficiency. Even in the demand equation, where the restrictions were *ad hoc,* the forecast errors were on average much smaller than under any other methodology. For the stock equation, in which the restrictions produced an identity, the results were even more pronounced.

Another result with practical applications is that, for some variables, very simple models requiring little in the way of data can provide relatively efficient forecasts, the best example being the effectiveness with which an ARIMA model forecasted CDEMD. The vector autoregression model also provided fairly efficient forecasts, especially as the forecast horizon lengthened.

Another related result is that it seems difficult to determine beforehand whether or not a method will work. Although Lamm reported success with an open autoregressive model at a more aggregated

Table 7—Mean absolute error (MAE) of CDEMD forecasts

Model	Forecasting period			
	1	2	3	4
Restricted equations original series	132.9	133.0	110.7	121.9
Restricted equations differenced series	31.9	41.9	48.0	48.2
ARIMA model	47.6	60.4	64.0	68.6
Vector autoregression	60.0	63.2	53.1	60.3
Open autoregression	332.9	310.9	254.9	250.1

Table 8—Mean absolute error (MAE) for CSTKS forecasts

Model	Forecasting period			
	1	2	3	4
*Restricted equations original series	132.9	133.0	110.7	121.9
*Restricted equations differenced series	31.9	41.9	48.0	48.2
ARIMA model	158.3	262.7	351.9	385.0
Vector autoregression	418.0	331.1	509.2	433.7
Open autoregression	332.5	597.0	587.5	718.5

*Because CSTKS is determined by an identity in the restricted equation models, the MAE here is equal to the error associated with CDEMD.

level, the model here grew quite large and, plagued with multicollinearity problems, failed to produce useful forecasts. Although methods with the potential power to reduce the problem are available, such applications here would have detracted from the model's value as a quick alternative method to restricted equation forecasts.

Conclusions

Well-constructed simultaneous restricted-equation models are useful and marketable policy tools, despite practical and theoretical problems. Because of their usefulness, econometric models, forecasts, and analysis based on models have become more abundant and more comprehensive. However, the types of summary statistics generally made available to model users provide limited information with which to evaluate the model or its forecasting capabilities. As has been shown here, modelers who attempt to maximize the summary statistics normally reported in economic journals need not arrive at an optimal model. Model users, judging two competing models on the basis of the summary statistics normally reported by economic journals, will not necessarily be able to choose the better model.

Information about the structure of the error terms is essential to accurately evaluate any model and its summary statistics. Such information is neither generally requested nor provided by most economic journals. Furthermore, such information is generally neither requested by buyers nor provided by sellers of models or model forecasts.

Another effective method for evaluating a model and its forecasting abilities is to compare the out-of-sample forecasts of complicated models with the out-of-sample forecasts of simple models. ARIMA and vector auto-regression models are two such models which provide relatively good forecasts.

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Research Review

New Developments in Macroeconomic Theory: A Prospectus and Appraisal

By Roger K. Conway and James R. Barth*

> Although the battles that are fought in... [the economics literature]...appear to be fought with antique pop guns, the bullets are real and they may soon be fired at you....
>
> Robert Solow

Introduction

There is no longer, if there ever was, any question but that the workings of the macroeconomy have major impacts on the agricultural sector. Indeed, one of the invited paper sessions at the 1982 American Agricultural Economics Association (AAEA) meetings focused exclusively on "Current Macroeconomic Policies and U.S. Agriculture." Furthermore, some of the important influences of macroeconomic policy on the farm sector have been recently and extensively discussed in an excellent survey article by Prentice and Schertz (*80*) and in an article by Tweeten (*108*) in which the latter goes so far as to state that "the economic structure of the farming industry in the long run will depend more on federal taxation, money supply and trade policies in the Federal Reserve System, the Internal Revenue Service, and Department of State than on commodity programs in the U.S. Department of Agriculture" (*108*, p. 864).[1] Although the Prentice and Schertz monograph provides a useful discussion of standard monetarist and Keynesian theories, both of which to varying degrees provided the foundation for the economic policies implemented during the fifties through the midseventies, a new theory—the new classical macroeconomics—also merits attention. This recent theory gained prominence during the turbulent seventies, when the claims that fiscal and monetary policy could smooth the business cycle were called into question. Broadly contemptuous of the fact that the predominant views of the macroeconomy seemed to be unable to explain the recent period of simultaneously high inflation and sluggish output growth, some economists claimed this new theory could explain how to reduce inflation quickly and without loss of output.

We believe it is in the interest of the agricultural community to become familiar with this new theoretical development in macroeconomics because it may, as Tweeten suggests, have a substantial influence eventually on what happens in the agricultural sector. Moreover, this theory already has operational components directly applicable to the agricultural sector (see, for example, *13, 33, 37, 41, 102, 105*). Our article will, therefore, attempt to present, as well as to appraise, the new classical macroeconomics. One of our major conclusions is that the new classical economics is far more complementary with traditional macroeconomic theory than either advocates or detractors have thus far emphasized.

The New Classical Macroeconomics

A crucial element of the new classical approach is the modeling of business cycles within an equilibrium framework. As it is assumed that expectations are formed rationally, these models are frequently referred to as rational expectations models. However, this assumption is but one of several identifying this class of models.[2] The other major assumptions are: (1) perfect wage and price flexibility will occur, leading to continuous market clearing; (2) economic agents possess imperfect information about some important economic variables; (3) some form of the "Lucas-Rapping supply function" applies in which (as described later) nominal shocks have only temporary influences on real variables; and (4) a variant of the natural rate hypothesis holds in which perceived nominal variables produce no real effects. The importance of these assumptions will become clear as we describe a basic rational expectations model.

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[1] Italicized numbers in parentheses refer to items in the References at the end of this article.

[2] We shall describe these models synonymously as new classical, rational expectations, or equilibrium cycle models.

Rational Expectations Defined

The concept of rational expectations was first postulated by Muth (*74*) when he suggested that, as expectations are informed predictions of future events, it seems only reasonable to consider them exactly like predictions from a relevant theory of economic behavior. This postulate can be stated more rigorously as follows. Based on a given information set, the subjective probability distributions of outcomes are assumed to be distributed about the objective probability distributions of outcomes. This equating of subjective and objective probability distributions is based on the notion that, because information is a relatively scarce commodity, economic agents within an economy will use it optimally. More specifically, the structure of the pertinent economic model is assumed to be information which is known and thus used in the formation of expectation.[3] However, it is not assumed that the predictions of individuals are always correct (perfect foresight) or that expectations are the same for all agents (homogeneous expectations).

Malliaris and Brock (*70*) contend that rational expectations is a plausible theory because: (1) the hypothesis is applicable to all dynamic problems, and expectations in different markets would not have to be treated in different ways; (2) if expectations were not moderately rational, there would be opportunities to make profits; and (3) rational expectations is a hypothesis that can be modified with its analytical methods remaining applicable in systems with incomplete or incorrect information (*70*, p. 153).

A Basic Rational Expectations Model: Its Policy Implications

We will use a simple model developed by Sargent and Wallace (*90*) to describe the basic features of a rational expectations model. We start by assuming that the demand for goods and services (an IS equation) is given by the following equation (where all subsequent variables except the interest rate are in natural logarithms):

$$Y_t^d = Y_p - \alpha(r_t - \dot{P}_t^e) + u_{1t} \qquad (1)$$

where:

Y_t^d = demand for goods and services;
Y_p = the trend or natural level of output;
r_t = the nominal rate of interest;
P_{t+1}^e = expected price for period t + 1 determined in period t;
$\dot{P}_t^e$ = The expected rate of inflation determined in period t − 1—that is, $(P_{t+1}^e - P_t^e)$; and
u_{1t} = a random error term such that $u_{1t} \sim N(0, \sigma_1^2)$.

Turning to the money market (or an LM equation), we can assume that the money demand function is given by:

$$M_t^d = P_t + \gamma Y_t - \tau r_t + u_{2t} \qquad (2)$$

where:

M_t^d = the demand of money,
P_t = actual price level during period t,
Y_t = equilibrium output, and
u_{2t} = a random error term with $u_{2t} \sim N(0, \sigma_2^2)$

whereas the money supply (M^s) is exogenously determined by monetary authorities and will be explained more fully.

Aggregate supply is assumed to be given by the following equation:[4]

$$Y_t^s = Y_p + \beta(P_t - P_t^e) + u_{3t} \qquad (3)$$

where:

Y_t^s = aggregate supply,
U_{3t} = a random error term such that $u_{3t} \sim N(0, \sigma_3^2)$,

and the other variables are as described above.

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[3] Adaptive expectations is another widely employed expectation scheme, whereby the expectations of a variable are based solely on past values of that variable. Only in the unlikely case that prices, when actually modeled, are found to display a random walk will rational and adaptive expectations yield the same results.

[4] Notice that when there is no uncertainty so that $P_t^e = P_t$, then the aggregate supply curve becomes vertical and one obtains essentially the classical model—hence, the term new classical macroeconomics. For the theoretical development of the new classical aggregate supply curve, see *(58, 60, 62, 94)*.

Using equations (1), (2), and (3) and the equilibrium or market clearing conditions, $M^d = M^s = M$ and $Y^d = Y^s = Y$, to solve for the reduced-form price equation, one obtains:

$$P_t = \frac{-\gamma\alpha}{D} Y_p + \frac{\tau\alpha}{D} P^e_{t+1} + \frac{(\beta - \alpha)\tau + \beta\gamma\alpha}{D} P^e_t + \frac{\alpha}{D}(M_t - u_{2t}) + \frac{\tau}{D} u_{1t} - \frac{\tau + \gamma\alpha}{D} u_{3t} \qquad (4)$$

where $D = \beta\tau + \alpha + \gamma\beta\alpha$.

Assuming that expectations are formed rationally, then:

$$P_t = E(P_t \mid \Omega_{t-1}) \qquad (5)$$

where Ω_{t-1} represents the information set available to every economic agent at time $t - 1$.

Applying the conditional expectations operator to equation (4), one obtains:

$$E(P_t) = \frac{-\gamma\alpha}{D} Y_p + \frac{\tau\alpha}{D} E(P_{t+1}) + \frac{(\beta - \alpha)\tau + \beta\gamma\alpha}{D} E(P_t) + \frac{\alpha}{D} E(M_t) \qquad (6)$$

This equation states that the conditional expectation of the price level is determined by future price expectations and the expected money stock. Next, we subtract the expected price level from the actual price level, yielding:

$$P_t - E(P_t) = \frac{\tau}{D} = u_{1t} - \frac{(\tau + \gamma\alpha)}{D} u_{3t} - \frac{\alpha}{D} u_{2t} + \frac{\alpha(M_t - E(M_t))}{D} \qquad (7)$$

Substituting equation (7) into equation (3) yields:

$$Y_t = Y_p + \frac{\alpha\beta}{D}(M_t - E(M_t)) + u_t \qquad (8)$$

$$\text{where } u_t = \frac{\beta\tau}{D} u_{1t} - \frac{\beta\alpha}{D} u_{2t} - \frac{\beta(\tau + \gamma\alpha)}{D} u_{3t}$$

According to equation (18), output depends only on the difference between the actual money supply and the expected money supply. To determine the expected money supply, we now postulate a policy feedback rule wherein policymakers attempt to stabilize output at the full employment level. This procedure is captured in the following equation:

$$M^s_t = M_{t-1} + \lambda(Y_p - Y_{t-1}) + u_{4t} \qquad (9)$$

where u_{4t} is a random error so that $u_{4t} \sim N(0, \sigma^2_4)$.[5] This equation indicates that the monetary authorities will increase (decrease) the money supply during recessionary (inflationary) periods. Based on the principle of rational expectations, one can now solve for the expected money supply. Doing so yields:

$$E(M_t \mid \Omega_{t-1}) = M_{t-1} + \lambda(Y_p - Y_{t-1}) \qquad (10)$$

Subtracting equation (10) from equation (9) yields $M_t - E(M_t) = u_{4t}$. This means that equation (8) reduces to:

$$Y_t = Y_p + \frac{\alpha\beta}{D} u_{4t} + u_t \text{ or}$$

$$Y_t - Y_p = \frac{\alpha\beta}{D} u_{4t} + u_t \qquad (11)$$

The importance of this result is that the systematic component λ is not present in equation (11), which means that the monetary authorities cannot systematically affect output. Only the random element in the money supply, u_{4t}, affects output, but this component represents unanticipated movements, and is thus uncontrollable by the monetary authorities. On the basis of this finding, Sargent and Wallace established the "superneutrality" proposition of monetary policy: differences between actual and full employment output follow a random walk, and thus are not subject to being stabilized by the monetary authorities.[6]

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[5] It is assumed throughout that $E(u_i u_j) = 0$ for all $i \neq j$ where $i,j = 1, \ldots 4$. This assumption is somewhat dubious as equations (1-3) are simultaneous so that joint probability distributions are implied. Sargent and Wallace make this assumption for computational tractability; however, as this need not be the case, this particular approach is limited.

Okun (*77*) and Schultze (*92*) criticize Sargent and Wallace for pretending that "all the world were a wheat pit." They argue that the real world is predominantly composed of customer markets, not auction markets, which are based on forward contracts, thereby implying considerable wage and price rigidities. Subsequent papers have, therefore, introduced rigidities into the Sargent-Wallace model and have then examined the resulting implications.

Whether or not the strong policy impotence result holds will depend mainly on the specification of output given by equation (3). In one study, when Lucas (*57*) assumes that production is withheld in anticipation of higher prices in future periods, actual output is then influenced by the systematic component of the policy feedback rule. In a subsequent article, however, McCallum (*65*) once again obtains the Sargent-Wallace neutrality result using the Lucas supply function specification, but only after discounting $E(P_{t+1})$ appropriately by the rate of interest.

The policy impotence hypothesis no longer holds once certain modifications are made in the output supply equation. For example, Fischer (*36*) specifies a model in which wages are constant for two periods because of contracts. In this case, the aggregate supply function becomes:

$$Y^s_t = \frac{1}{2} \sum_{i=1}^{2} (P_t - \underset{t-1}{E} (P_t)) + u_{3t} \qquad (12)$$

Aggregate supply is now no longer influenced only by the current informational error between actual and expected price. Instead, monetary policy, even with the assumption of rational expectations, can systematically influence real wage rates and thereby aggregate output because the authorities' actions will not be fully and contemporaneously reflected in current nominal wage rates. Taylor (*101*) and Gray (*46*) also develop similar models of overlapping wage contracts in which, in any given period, some contracts are being negotiated while others are still in force. In this situation, monetary policy is again effective because of the sluggishness of wages relative to prices.

Phelps and Taylor (*79*) specify yet another model, one based on the assumption that firms set their prices in the period previous to the one in which they sell their output. This type of advanced price setting enables monetary policy to affect real variables without influencing prices in the current period. The following aggregate output equation as posited by Phelps and Taylor incorporates this price stickiness:

$$Y^s_t = \alpha_0 + \alpha_1 E_{t-1}(P_t) + \alpha_2 P_{t-1} + \alpha_3 M_t + u_{3t} \qquad (13)$$

In this type of model, monetary policy can clearly affect real output. However, McCallum (*67*) demonstrates that slow price adjustment *per se* is not incompatible with the policy ineffectiveness hypothesis. One can demonstrate this hypothesis by positing an aggregate supply equation in which the superneutrality result holds no matter what rigidities are introduced. The key factor guaranteeing the policy ineffectiveness hypothesis in this situation is that the length of the period during which the wage is rigid is no longer than the period for policy action to be effective.[7]

The Problem of Persistence

Strictly interpreted, the expectational errors described above should be randomly distributed over time if expectations are formed rationally. If so, it directly follows that output should be uncorrelated over time. Needless to say, this is not the case when one examines the data. Quite the contrary, all the available evidence indicates that output and unemployment are highly serially correlated. The observed "persistence" of positive output-price and money-output relationships over business cycles would seem to contradict policy neutrality.

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[6] Note that all right-hand side variables are considered orthoganal to the error term when rational expectations models are estimated. Thus, tests of causality or, more properly, exogeneity are closely linked to the time series models of rational expectations theorists (see, for example, (*84*, *89*)). Unfortunately, causality tests are subject to serious theoretical and empirical complications which, in our view, limit their usefulness. See (*28*) for a discussion of this point. Note also that the emphasis is on cyclical movements, which is why applied rational expectationists detrend and deseasonalize their data. One attempts to work with economic time series which are, in other words, covariance stationary and purely indeterministic.

[7] Readers are advised to consult the survey article by McCallum (*68*) for a more complete discussion of alternative supply functions and their implications for policy effectiveness. One might add that, although McCallum's counterexamples are effective debating points, they appear somewhat artificial in describing real world relative rigidities between wages and policy implementation. Perhaps recognizing this, McCallum (*66*) provides a more recent and shrewder critique of sticky wage/price models.

Most rational expectations supply functions are modified to include a lagged output term as an explanatory variable to account for this observed serial correlation. However, this procedure for dealing with persistence has been criticized by Modigliani (*73*) as being *ad hoc*, as no rigorous theoretic framework is provided to justify the inclusion of this additional variable. In response, rational expectations adherents have subsequently offered several explanations to account for the lagged term. One explanation is based on the notion that costs of adjustment prevent instantaneous adjustments. Another, advanced by Lucas (*55*) and explicitly derived by Sargent (*88*), is based on information lags which lead to serial correlation. More specifically, a model based on Phelps' (*78*) work with dispersed markets, called "islands," and informational discrepancies is used to generate demand-induced price-employment correlations in the economy as well as to provide a theory of the persistence effects associated with aggregate-demand shocks. Lucas' basic argument is that economic agents engaged in market activity are exposed to a continuum of dynamic and unanticipated opportunities that will not remain constant while an individual conducts an information search. In such a situation, real and nominal effects cannot be immediately disentangled so that serial correlation occurs.

Another explanation for serial correlation or persistence is based on the existence of inventories. Blinder and Fischer (*14*) demonstrate that a sudden price change can be met by altering inventories so that changes in production will be less than sales. If prices remain the same in the next period, production could be increased to rebuild inventories to their original level. As a result, a lagged increase in production of varying duration would occur while the original inventory stock is being restored. Even if no surprise price change occurred in time t + 1, the output gap in time t + 1 would still depend positively on the output gap in previous periods. Unfortunately, none of these explanations for persistence appears entirely satisfactory.

Rational Expectations, Stability, and Multiple Equilibria

Thus far we have examined the rational expectations hypothesis as well as the use of lagged variables to explain observed price and output movements. Whenever lagged variables appear, however, the question of convergence of the economic system naturally arises. When considering this question, Shiller (*95*) contends that "these (rational expectations) models may explode rather than converge. . . because the expectations mechanism held at time t influences the behavior of Y and hence the revised expectations mechanism in time t + 1. Because of the arbitrary nature of the adjustment mechanism, we cannot generally tell which models will explode and which will not." Furthermore, he argues that the possibility exists that "we are left with a fundamental indeterminancy for the solution of rational expectation models and infinity of potential solutions for all but those degenerate models which yield zero order difference equations."

Both these issues are serious because they suggest that rational expectation models may be specified so as to permit the possibility of self-fulfilling unstable expectations. As Flood and Garber (*38*) note, a price bubble can develop when the actual market price depends positively on its own expected rate of change. Because economic agents under rational expectations will not make systematic prediction errors, price and its expected rate of change are positively related, suggesting a similar relationship between price and its actual rate of change. Thus, one may be in a situation in which an "arbitrary self-fulfilling expectation of price change may drive actual price changes independently of market fundamentals" (*38*, p. 746).

Rational expectations proponents vary in their response to this criticism. Muth (*74*) originally attempted to dismiss this issue by pointing out that "for a bounded solution, the coefficient of the larger (unstable) root vanishes; the initial condition is then fitted to the coefficient of the smaller root." In other words, if the general solution for price is:

$$P(t) = A_1\lambda_1 + A_2\lambda_2 \tag{14}$$

and $\lambda_2 > 1$, then $A_2 = 0$ and A_1 is determined by the initial condition. However, Sargent and Wallace

(*90*) argue against unstable roots because of "a terminal condition that has the effect of ruling out 'speculative bubbles.' " Minford (*71*) expresses perhaps the most popular defense against this particular weakness when he asserts that past history has not provided examples of prices exploding to infinity. Because this information becomes a part of the model itself, it is presumed that one need only consider converging solution.[8] Nevertheless, those who believe the dynamics for explosive behavior still exist have challenged this argument, although admitting that some exogenous action (for example, policy) may serve to dampen the instability.

Burmeister (*20, 21*) believes the non-uniqueness question may be more serious for equilibrium models than for those models wielding "free parameters" and producing disequilibrium price adjustments which are anathema to the new classicists. In any event, the issues of stability and multiple equilibrium have not yet been adequately resolved in the new classical economics literature.

The Lucas Critique and Time-Inconsistency

Perhaps the most important consequence of the rational expectations literature is hypothesis, formulated by Lucas (*56*), that the behavior of economic entities is not invariant to changes in the economic environment. For example, when the Government changes its policy rules, it is assumed that individuals will observe this change and modify their behavior so as to maximize their objective functions. This means that traditional econometric models may be seriously flawed when used for policy analysis because these models simply extrapolate past behavior based on old policy rules into the future. In econometric parlance, the linkage between behavioral and policy coefficients (referred to as "cross-equation restrictions") needs to be taken into account. Stated another way, the coefficients in those equations describing the behavior of economic agents (such as supply and demand fuctions) should be restricted by the coefficients (or policy parameters) in those equations modeling economic policies. Hansen and Sargent (*48, 49*) are among the first to do empirical work taking these restrictions into account, particularly with respect to large econometric models. Their analysis suggests that incorporating the appropriate the appropriate restrictions will be extremely difficult.[9]

Kydland and Prescott (*53*) and Prescott (*81*) have taken the "Lucas critique" one step further by arguing that formulating Government policy using optimal control theory is invalid when economic agents maximize their objective functions using rational expectations because any policy result will be either "time-inconsistent" or suboptimal. By time-consistency, they mean the following. Suppose individuals formulate their optimal behavior in an initial period t where they have accurately anticipated all future Government policy variable magnitudes for each time period over a finite horizon. If these individuals follow their optimal plan until a period j and then recalculate an optimal plan for all succeeding periods, the original plan is considered to be time-consistent when the recalculated optimal plan is simply the continuation of the original plan derived in the first period for all periods after j. Kydland and Prescott further state that when expectations are formed rationally by economic agents and when they are aware of the authorities' decision rule, the Bellman principle of optimality is inapplicable. To understand their point, consider the optimal rule for patents as an instance of time-inconsistency. Ideally, the Government would like to break its commitment with firms enjoying monopoly patent rights and thereby have the advantage of a competitive environment with the current technology now used under existing patents. However, the Government would also have to pledge to honor future patents to stimulate new inventions. In a future time period, however, these new inventions would be old and, once again, the Government could terminate all past patents. As expressed by Barro, "the policy that optimizes at all dates subject to the initial conditions applying at those dates is subject to a time-inconsistent property where earlier plans are not followed through" (*4*, p. 59). Once anticipated by private individuals, a steady or fixed policy is a time-consistent policy, but it results in suboptimal technological advance. Analogous policy considerations can easily be imaged within the context of the agricultural sector, such as international trade agreements, buffer stock price stabilization programs, and crop set-aside programs. Some critics, like Modigliani (*73*), have argued that time-inconsistency would emerge as an obstacle only when policymakers and private individuals maximize heterogeneous and incompatible objective functions. However, Calvo (*22*) has provided a counter example, demonstrating the possibility of a time-inconsistent result emerging *even* when policymakers and private individuals have the same objective functions.

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[8] However, price bubbles have occurred with individual commodities whether for 17th century Dutch tulips or for world sugar during the 1973-74 commodity boom. See Conway (*25*) for a discussion of this issue.

[9] The insight of Lucas is independent of rational expectations so that one could conceive of "irrational expectations" producing the same parameter instability in econometric models. See (*12*) for estimation of a policy reaction function.

Kydland and Prescott's finding apparently supports the superiority of rules rather than discretion in the conduct of economic policy. The longstanding debate between rules and discretion can thus be revised in the framework of the new classical models to distinguish between rules without feedback and rules with feedback. Now that "optimal" policy rules might be suboptimal or time-inconsistent, a dilemma emerges as to whether policymakers should adhere to their original optimal rule or subsequently revise it. Kydland and Prescott prefer adherence, whereas Lucas and Sargent (*64*) believe this rule does not necessarily follow, based on the assumptions of their model. Instead, Lucas and Sargent contend that it is entirely possible that policymakers could achieve a higher level of total (additive) intertemporal utility by continually revising their plans every period than they could by adhering to an initial policy program. The counterargument in favor of rules put forth by Kydland and Prescott is that the Government runs the risk of "crying wolf." They specifically claim that policy announcements by the authorities would lose their credibility if the announcements were continually subject to revision. Lack of credibility, in turn, would increase the variance of an individual's reaction function, make its identification more problematic, and thus lead to indeterminate consequences. To avoid this situation, Kydland and Prescott, as well as Lucas, suggest that one simulate over a number of different policy rules and then implement the one with the superior operating characteristics. The significance of this proposal is that authorities would no longer seek the *optimum optimorum*, but would be forced to settle for a second best solution.

Credibility: Can Disinflation be Painless?

According to rational expectation proponents, credibility by both monetary and fiscal authorities is necessary for stable economic growth. In particular, inflation can be reduced only by a change in the current and prospective fiscal and monetary policy regimes that will reduce inflationary expectations. Temporary restrictive fiscal and monetary actions will not suffice; only a permanent change in the policy regime from "what is perceived as a long-term Government policy involving the high average rates of government deficits and monetary expansion in the future" will keep (*87*, p. 2). If such an "abrupt" change in Government strategy "is sufficiently binding as to be widely believed," Sargent argues that the cost in foregone output and unemployment could be small. As evidence for his position, Sargent (*86, 87*) cites events in Austria, Hungary, Germany, Poland, France, and Czechoslovakia during the twenties as examples where "credible" policies ended dramatic hyperinflations or prevented them from occurring at near-zero output cost.

Imposing a gold standard on a government is one way to establish credibility, Sargent contends, because the government issuing demand notes and long-term debt would promise to convert any outstanding notes into gold, on demand, or to levy sufficient taxes to honor its debt. That government would be expected to meet its debt obligations while simultaneously refraining from *seigniorage.*[10]

Sargent and Wallace (*91*) and Sargent (*85*) rely on the credibility issue to argue for coordination of monetary and fiscal policies because of their connection via the Government budget restraint. The Federal Reserve Board's policy regarding open market operations will influence the flow of potential *seigniorage* revenues which, as noted below, is a part of the intertemporal budget constraining fiscal policy. They show that this situation strongly influences the desirability of Friedman's "k-percent rule" for the monetary base. Thus, one can envision a situation where such a rule is undesirable, whether over the long run or not. First, it could happen when the amount of outstanding Government interest-bearing debt is so large that a pretax real interest rate greater than the rate of growth in fiscal output is needed to place the debt. Second, a situation could occur in which the monetary and fiscal authorities clash. If fiscal policy dominates monetary policy so that there is an ironclad rule for Government expenditures and explicit taxes, then a concomitant monetary accommodation may be needed to finance the growth rate of total Government indebtedness. Finally, there is the case in which the deficit path the authorities select is significantly in deficit in a present value sense. This condition seems to represent the economic situation in many Latin American countries today.

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[10] *Seigniorage* is revenue from money creation. For an analysis, see (*10*).

Arguments in support of the Friedman k-percent rule hinge on whether or not the dominating fiscal regime just described occurs. A k-percent rule doggedly adhered to by the monetary authorities is a way to discipline fiscal policy. Thus, a monetary authority that is stronger than its fiscal policy counterpart will announce and administer a k-percent rule for the monetary base, thereby dictating the flow of *seigniorage* revenues for the fiscal authorities, who are then forced to lower Government expenditures, increase tax revenues, and/or employ bond financing so as to balance the budget. This scenario seems to describe the current situation in West Germany where the Bundesbank has adopted such a rule and appears to have successfully adhered to it.

New Classicism: Empirical Results

Econometric investigations of the rational expectations hypothesis, although numerous, suffer from severe limitations. One problem is that rational expectations and adaptive expectations are "observationally equivalent," as Sargent (*89*) demonstrated. This means that both expectational theories may produce results that are equally consistent with the data. As a result, economists cannot test directly which theory is best supported by the evidence.

Indirect tests have therefore been attempted, but with mixed results. Barro (*7, 8, 9*), in particular, has produced many studies attempting to determine whether only unanticipated money (or money shocks) affects output. His findings indicate that anticipated money growth is neutral, whereas unanticipated money growth has real effects. However, by respecifying the way expectations are formed, Small (*96*) and Gordon (*42*) generated results that contradict Barro's findings. This reversal of results points out a weakness in testing rational expectations models; namely, that there is a joint hypothesis involving the specification of the model as well as the expectations mechanism.

Furthermore, only if strong assumptions hold can parameters be found that remain invariant under policy changes (as Lucas desires). In this situation, a varying parameter approach is the most general because it makes fewer *a priori* assumptions about the economic environment. If the coefficients do vary, then it is impossible "to roll back econometric models to the levels where parameters do remain invariant" (*106*, p. 12). Swamy, Barth, and Tinsley (*99*) have also demonstrated that there are serious problems with the ARMA models used to estimate the unobservable exogenous processes in the new classical models. They point out specifically that stationarity has been inappropriately assumed and that there are identification problems associated with stationary models and unrestricted reduced forms which have not yet been addressed.[11]

One very suggestive study by Neftci and Sargent (*75*) argues that a change in the distributed lag relationship of output on the actual money supply will occur when the feedback policy rule of the monetary authorities changes at some point in time if the neutrality hypothesis is true, whereas no change will occur if a relationship between output and innovations in the money supply is found. They hypothesized a policy break during 1964 and, using a Chow test, found the evidence supported the neutrality proposition. Yet the Chow test is not robust as there is a difficulty in correctly characterizing policy behavior by time periods. Furthermore, Neftci and Sargent's alternative hypothesis for this test is only a single arbitrary break in policy regime. In fact, a structural break is possible in every period and, on that basis, Swamy and Tinsley (*100*) show that the Chow test is statistically inconsistent, and Resler, Barth, Swamy, and Davis (*82*) discuss problems with detecting structural change based on existing tests.

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[11] Conway and Fryar (*27*) provide a further discussion of econometric estimation difficulties with rational expectations models.

Theoretical Weaknesses of the Rational Expectations Approach

One fundamental problem analyzed by Swamy, Barth, and Tinsley (*99*) is the contradiction between the Muth-Lucas-Sargent rational expectations hypothesis and the Bayesian definition of rationality based on coherence and other consistency requirements.[12] As they point out, the problem is that the rational expectations hypothesis requires the equality of subjective and objective probabilities (they rely on three different definitions of objective probabilities for comprehensiveness) of events. In contrast, the subjective Bayesian principle of rationality relies entirely on subjective probabilities, thereby requiring no such equality. They also argue that the rational expectations hypothesis may be considered a linguistic axiom, this being too restrictive for practical use. As an alternative, Swamy and others build a competitive, logically consistent, microeconomic model based on a Bayesian framework of subjective expectations which, when aggregated across individuals, becomes a stochastic coefficient macroeconomic model suitable for empirical use.

Rational expectations proponents are circumspect in discussing how the "true" underlying structure of the economy becomes known, how fast agents can obtain information about this structure, and what the information costs are. The assumption by proponents appears traditionally to be that individuals acquire knowledge of the structure of the economy rapidly and freely for epistemological ease. A more realistic view is that it takes a long time for economic agents to discover (if they ever do) the true structure of the economy and hence to adjust their expectations appropriately. For this reason, rational expectations may be more appropriately considered to be a longrun, steady state phenomenon. However, if a break in structure occurs during the transition period toward a longrun equilibrium, problems arise. A more reasonable assumption, noted by Friedman (*39*), is that knowledge of the new structure is costly to obtain and, therefore, will be acquired adaptively. If so, activist policies under this regime will generate real effects. Using Hicks' (*50*) terminology, real time is abstracted from the new classical models, and logical, not historical, time becomes the criterion, as rational expectations proponents consider the length of temporal expectations formation to be "whatever it takes" to allow individuals the opportunity to acquire sufficient information to assure that expectations are fulfilled.

Finally, a contradiction to the Sargent-Wallace policy impotence result described above arises when one takes into account the portfolio balance theory as put forth by Buiter (*19*), Tobin (*103*), and Hadjimichalakis (*47*). Perfectly anticipated increases in the money growth rate will increase the perfectly anticipated inflation rate and thereby induce substitution of capital for real balances. This increases the capital stock and hence generates impacts on output and employment so that money is not neutral after all.

Implications of the New Classical Macroeconomics for the Government Budget

The new classical macroeconomics provides interesting policy prescriptions when one considers the Government budget. An examination of Christ's (*24*) formulation of the Government budget restraint (GBR) clarifies this issue. As Christ noted, the Federal Government must satisfy the following constraint:

$$G + iB = T + dM + dB - dA \tag{15}$$

or:

$$\text{Deficit} = G + iB - T = dM + dB - dA \tag{16}$$

where i is the nominal rate of interest ($i = r + \pi^e$), π^e is the expected inflation rate, and r is the real rate. This equation states that Government expenditures on goods and services (G) minus taxes (T) plus interest paid on the Government debt outstanding (iB) must be financed by issuing more outside money (M) and/or bonds (B) and/or selling assets (A). Equation 15 clearly shows that the Government does not have a free hand to determine all the macroeconomic policy variables simultaneously. Instead, at least one policy tool must be endogenously determined. Another point made clear by equation (15) is that a government's budget constraint is not quite the same as an individual's. Notice specifically that by increasing the money supply and generating inflation, the Government can reduce its real deficit. This it does by reducing the real value of Government bonds and other outstanding liabilities. However, because the nominal interest rate incorporates inflationary expectations, this inflationary tax may be moderated to the extent the inflation is fully anticipated and is thereby embodied in the rate of interest (that is, $\pi = \pi^e$). The specific influence of inflation on the GBR is shown by the following equation.[13]

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[12] See (*31*, chapters 6 and 7) for a rigorous definition of these axiomatic principles. Coherence, as defined by de Finetti (*30*), is the most important consistency requirement. It amounts to restricting a given belief system so that it would be impossible for any individual to lay multiple fair bets in such a way that the individual wins under all possible outcomes. When coupled with other axioms, the suggestion is that the probabilities of elementary events will sum to 1 and that conjunctive and disjunctive events will obey both the product and addition rules.

$$G/P + rB/P + \pi M/P - T/P - \pi A/P = d(B/P) + d(M/P) - d(A/P) \qquad (17)$$

The new classical school believes that the influence of tax cuts on inflation and output depends on whether or not the deficit is bond-financed. If it is, there will be neither inflation nor real effects if individuals engage in "ultrarational" behavior.[14] In this case, as Barro (*1, 6*), David and Scadding (*29*), and others argue, there is no essential difference between debt and tax finance. This view is also referred to as the "Ricardian equivalence" theorem because it traces back to Ricardo's tract on the full capitalization of the tax when bond sales are used to finance the debt. A Government sale of securities which is fully anticipated by individuals means that private spending will be reduced by exactly the same amount as the bond-financed deficit because individuals will realize that the issuance of Government securities means a higher future tax liability for which savings must be increased. Under this behavorial assumption, aggregate demand remains constant when taxes are cut, leaving prices and real output unchanged. Moreover, the IS-LM curves remain unchanged, thus leaving the interest rate unaffected. However, the composition of output has changed; private spending has dropped by an amount equal to the increase in Government spending.

Of course, strong assumptions underlie this analysis. In particular, the model depends on finite-lived individuals linked across generations by bequests.[15] Because of this assumption and other factors, many individuals have criticized these models, including Feldstein (*35*), Buchanan (*17*), Buiter (*18*), Drazen (*32*), Carmichael (*23*), Stokey (*98*), Brewley (*16*), and Tobin (*104*). The main criticisms are: (1) altruism may be absent between parent and child; (2) the generational chain may be broken by people without children; (3) constraints on wealth transfers among generations may arise from imperfect rental, mortgage, and annuity markets; and (4) survival is uncertain so that the private risk of death will be greater than the social risk. Empirical evidence on the ultrarationality hypothesis, as discussed in a survey article by Stevens (*97*) and a recent study by Feldstein (*34*), seems to indicate that discounting of bond-financed deficits is less than complete.

In contrast to taxes, the output effects of Government expenditures, as demonstrated by Barro (*5*), depend on whether they are perceived as temporary or permanent. In the case of permanent Government expenditure changes, the assumption is made that tax rates will be adjusted so as to generate sufficient revenues to match the change in permanent Government expenditures. Thus, concomitant changes in tax rates will only result from a new and permanent level of Government expenditure. Based on this reasoning, Barro (*2, 5*) and Kydland and Prescott (*52*) argue that if all changes in Government spending were permanent, then tax rates should be constant or nearly constant over the business cycle so that the budget would be balanced on average, with surpluses and deficits arising only from random output fluctuations.

In contrast, according to Barro (*3*), temporary tax and Government expenditure changes significantly influence output because incentives for intertemporal substitution of work and production are created. The greatest incentives arise in the case of transitory expenditures which are not considered as close substitutes for private spending (for example, wartime expenditures). Individuals recognize that income is temporarily increased and will react by temporarily increasing their labor supply. At the theoretical level, the basic result is that aggregate demand is unaffected while aggregate supply increases. As far as empirical work is concerned, Barro (*5*) found support for the view that temporary movements in defense purchases, generally associated with war episodes, almost doubled the response in output as compared with an equal and permanent shift in defense purchases.

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[13] Barth and Cordes (*10*), Barth and Morrell (*11*), and Turnovsky (*107*) discuss these and other dynamic considerations of the GBR.

[14] This behavior is used to explain "Denison's Law"—the observed stable relationship between gross private saving and gross national product in the United States.

[15] A recent and growing interest has developed in extending Samuelson's (*83*) overlapping generations model to examine a number of macroeconomic issues involving the intertemporal allocation of resources. We point this out because this type of model is an equilibrium one and assumes perfect foresight or rational expectations. See McCallum (*69*) and Kareken and Wallace (*51*) for a discussion of this literature.

Conclusions

A remarkably persuasive recent article by Boland (*15*) puts forth the proposition that an empirical test can be "true" only if the theory behind it has logical validity. More than 2,300 years ago, Aristotle constructed necessary conditions in the form of axioms for admissible arguments (that is, the principles of deductive reasoning). An argument is logical when two rules of inference hold: (1) *modus ponens*, which states that the truth of every assumption implies the truth of all conclusions, and (2) *modus tollens*, which states that when one conclusion is false, at least one assumption is false. It is unfortunate that discerning the truth of assumptions in economic theory is impossible because the problem of induction has never been solved. This problem means that the new classicists are unable to have their propositions logically verified from the truth of economic particulars.

As a result, the sophisticated falsification criterion developed by Popper and emphasized in economics by Friedman[16] considers the absolute truth of competing theories less important than testing them to discover which one has the most successful logically derived conclusions or predictions. This process may be performed by use of statistical techniques and historical evaluation. The new classicists have thus far succeeded in demonstrating observational equivalence with traditional models as well as in providing historical examples that seem to support some of their arguments. Yet others, like Gordon (*43, 45*), have produced historical counter examples supporting the view that a credible reduction of inflation by the authorities will be painful both in terms of unemployment and lost output.

The contribution of this new economic program may ultimately lie not so much in explaining the historical record as in ascertaining the microeconomic foundations used to produce its macroeconomic conclusions. The primary value of the new classical school may be its attempt to develop a coherent, economic explanation of subjective expectation formation by use of optimization and game-theoretic techniques so as to explain cyclical movements in key macroeconomic variables better. In this respect, the Lucas critique and the time inconsistency proposition have already emerged as potentially important contributions to interpreting and to reformulating econometric policy analysis.[17] Economists are now more aware that authorities are not playing a game against nature whose response is invariant to policy changes. Instead, the game is played against alert and rational agents who will react to policy stimuli. As a result, the new classical macroeconomics has been valuable in demonstrating how the length of real policy influence could be briefer and that the impact and predictability of those policies may be weaker than previously believed. Rather than viewing the new classical approach as a distinct macroeconomic program, we believe it may be more appropriate to consider it a part of the most important new scientific research program to emerge over the past decade: information theory and probabilistic economics.[18] This important research is being rapidly incorporated into all areas of economics, including public finance, industrial organization, labor economics, and public choice. Although the assumptions of new classical models tend to coincide with the underlying monetarist models, they can also be modified to yield disequilibrium Keynesian results.[19] The dramatic policy neutrality conclusion that resulted in such a highly controversial introduction of the new classical models is being toned down considerably by its originators. As Sargent and Lucas now state, "the point of this example was to show that within precisely that model used to rationalize reactive monetary policies, such policies could be shown to be of no value. It hardly follows that *all* policy is ineffective in *all* contexts" (*63*, p. 60) and "[t]his is a narrow special case, requiring arbitrary assumptions about information flow which has found little or no empirical support" (*93*, p. 11). This important qualification has only recently been admitted openly.

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[16] See (*26*) for a brief discussion of Popper, Friedman, and Lakatos, as well as other influences on economic methodology.

[17] We believe recent challenges to the Lucas critique by proponents of Vector Autoregression (VAR) models are misplaced. In addition to McCallum's (*66*) rebuttal to their argument, time series models such as VAR suffer from an incomplete theoretical development. See (*28*) and (*100*) for a more complete explanation.

[18] See (*54*) for an illuminating discussion of this research program.

[19] Notice that deviations of variables from their trend values may not be a sign of market failure, and furthermore, even if they are, the most appropriate way for the Government to intervene, if at all, may be through providing of information. One way is through the development of contingent claims markets, such as options and futures markets. See (*40*) and (*76*) for a full discussion.

In summary, we contend that research based on the new classical macroeconomics approach may be viewed as extending, rather than replacing, traditional models. Greater recognition of the stochastic nature of macroeconomics can only add to policymakers' understanding of how the economy works and can thereby lead to improved policy recommendations. Lucas (*59*) perhaps put these new contributions in the proper context: "[t]here is no point in letting tentative and, I hope, promising first steps harden into positions that must be defended at all costs" (*59*, p. 713).

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The Food and Agricultural Policy Simulator: An Addendum

By Larry Salathe, J. Michael Price, and Kenneth Gadson*

In a recent article, "On the Misuse of Theil's Inequality Coefficient," by Leuthold in this journal (*3*), the author criticizes the use of a version of Theil's inequality coefficient to validate the performance of an econometric model.[1] Referring to the article, "The Food and Agricultural Policy Simulator," by Salathe, Price, and Gadson (4), Leuthold suggests that "The [Theil inequality coefficients] U's reported by Salathe, Price, and Gadson are probably much too low, reflecting greater accuracy than is really the case." Leuthold also raises a number of issues regarding the use of the Theil inequality coefficient presented by Salathe, Price, and Gadson, and he indicates they should report additional validation statistics for the Food and Agricultural Policy Simulator (FAPSIM). The purpose of this note is to address some of the issues Leuthold raises and to present additional validation statistics for FAPSIM.

Three versions of the Theil inequality coefficient are frequently used to evaluate econometric models. These statistics are defined as follows:

$$U = (\sum_{t=1}^{T} (Y_t - \hat{Y}_t)^2)^{1/2} / [(\sum_{t=1}^{T} Y_t^2)^{1/2} + (\sum_{t=1}^{T} \hat{Y}_t^2)^{1/2}] \qquad (1)$$

$$U_1 = (\sum_{t=1}^{T} (Y_t - Y_t)^2)^{1/2} / [(\sum_{t=1}^{T} Y_t - Y_{t-1})^2)^{1/2} + \qquad (2)$$

$$(\sum_{t=1}^{T} (\hat{Y}_t - Y_{t-1})^2)^{1/2}]$$

$$U_2 = (\sum_{t=1}^{T} (Y_t - \hat{Y}_t)^2)^{1/2} / (\sum_{t=1}^{T} (Y_t - Y_{t-1})^2)^{1/2} \qquad (3)$$

where Y_t is the actual value of the variable at time t, $\hat{Y}_t$ is the predicted value at time t, and T is the number of time periods used for validation. The version used by Salathe, Price, and Gadson was the Theil U statistic.

Leuthold criticizes the use of this statistic for two reasons. First, this statistic is sensitive to additive transformations of the data as pointed out by Theil (*6*). In other words, if c denotes an arbitrary constant, then in general:

$$(\sum_{t=1}^{T} (Y_t - \hat{Y}_t)^2)^{1/2} / [(\sum_{t=1}^{T} Y_t^2)^{1/2} + (\sum_{t=1}^{T} \hat{Y}_t^2)^{1/2}] \neq$$

$$(\sum_{t=1}^{T} ((Y_t + c) - (\hat{Y}_t + c)^2))^{1/2} / [(\sum_{t=1}^{T} (Y_t + c)^2)^{1/2} +$$

$$(\sum_{t=1}^{T} (\hat{Y}_t + c)^2)^{1/2}]$$[2]

Leuthold suggests that this property is a major cause for concern, stating that, "if one moves the decimal point to the right for each variable, one can generate lower [Theil U] coefficients" (*3*). However, the type of transformation described by Leuthold is *not* an *additive* transformation, but a *multiplicative* transformation of the data. The Theil U statistic is not sensitive to multiplicate transformations because:

$$(\sum_{t=1}^{T} (\hat{Y}_t - Y_t)^2)^{1/2} / [(\sum_{t=1}^{T} Y_t^2)^{1/2} + (\sum_{t=1}^{T} \hat{Y}_t^2)^{1/2}] =$$

$$(\sum_{t=1}^{T} (cY_t - c\hat{Y}_t)^2)^{1/2} / [(\sum_{t=1}^{T} (cY_t)^2)^{1/2} + (\sum_{t=1}^{T} (c\hat{Y}_t)^2)^{1/2}]$$

for any nonzero constant c. This means that the units used to measure Y_t will not affect the value of the Theil U statistic. Therefore, the results presented in (*4*) do not reflect a "unit problem" as Leuthold suggests.

One also has to wonder why it is necessarily inappropriate to use a validation statistic which is not invariant with respect to an additive transformation, especially when this statistic is used to measure the predictive power of a model in terms of a price or a quantity series. For example, suppose we have a data series $\{Y_t\}$ for the price of some commodity, and suppose c is an arbitrary constant. Is there some meaningful interpretation that can be attached to the new series $\{Y_t + c\}$, and, if not, why would anyone be interested in determining how well the model predicts this variable?

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[1] Italicized numbers in parentheses refer to items in the References at the end of this note.

[2] Kost (*1*) claims that an additive transformation of this type will always reduce the value of U. However, this is not necessarily the case.

A second criticism of the Theil U statistic which Leuthold raises is that this statistic is bounded from above by 1. Leuthold seems to suggest that this property somehow leads to "low [Theil] U coefficients," thereby overstating the predictive power of a simulation model. As Theil showed, both the Theil U and U_1 statistics are constrained to lie between zero and 1. Both statistics will attain a value of 1 only if the predicted values represent the worst possible forecasts of the actual values (in the sense defined by Theil). Both the Theil U and U_1 statistics will equal 0 only if the model predicts the historical values perfectly over the entire validation period. Thus, both Theil's U and U_1 statistics are constrained to lie within the unit interval, and the end points of the interval may be interpreted as corresponding to opposite extremes in model forecast accuracy. The fact that Theil's U statistic is bounded from above by 1 is a useful property because the value of 1 provides a benchmark representing the worst possible model forecasts. This property in no way causes this statistic to overstate the predictive ability of a simulation model.

Leuthold also criticizes the use of Theil's U statistic because: " A U in actual values will always be less than a U when data are measured as changes." This statement implies that U will always be less than U_1 and U_2. However, using the example below, we can show that U is not necessarily less than U_1 or U_2. Suppose we have the following data series for Y_t and $\hat{Y}_t$:

Variable	Time period				
	0	+1	+2	+3	+4
Y_t	0.000	0.000	0.081	0.052	0.001
$\hat{Y}_t$	—	.000	.000	.000	.000

— = Not applicable.

then $U = 1$, $U_1 = 0.49045$, and $U_2 = 0.96246$.[3] Thus, there is no reason to expect that the Theil U statistic will always be less than either U_1 or U_2.

For all the data series included in the model validation of FAPSIM, however, the Theil U statistics are less than their corresponding Theil U_1 and U_2 statistics (see following table). But the Theil U_1 and U_2 statistics do not appear unreasonable, thereby confirming the earlier model validation results reported in (*4*). The Theil U_1 and U_2 statistics do result in changes in relative results among commodities as compared with the Theil U statistic as Leuthold suggests. This result is not surprising because of the differences in the properties of the Theil U, U_1, and U_2 statistics.

As Kost points out, each of Theil's inequality coefficients has different properties and interpretations (*2*). This means that each coefficient provides different information about the forecast accuracy of a model. The same is true of other statistics that have been proposed for model validation (*1*). No single measure provides a complete summary of a model's forecast ability as each has specific limitations. For this reason the mean absolute relative error, the turning point error, and the plots of actual versus forecasted values were presented in conjunction with the Theil U statistic in the validation of FAPSIM (*4*). Leuthold appears to be overly concerned with the properties of the Theil U statistic, failing to recognize the shortcomings of other validation statistics (such as Theil's U_1 and U_2 statistics). He also seems to have overlooked the additional information beyond the Theil U statistic presented in the validation of FAPSIM.

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[3] Although the values chosen for Y_t and $\hat{Y}_t$ in this example may appear artificial, they were, in fact, obtained from a validation study of a dairy-sector model (*5*).

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Theil inequality coefficients for the Food and Agricultural Policy Simulator

Variable description	Validation statistic		
	Theil U	Theil U_1	Theil U_2
Pork production	0.024	0.282	0.562
Beef production	.019	.347	.714
Broiler production	.014	.243	.495
Turkey production	.021	.300	.560
Egg production	.013	.552	1.157
Milk production	.006	.247	.516
Price of barrows and gilts	.048	.272	.512
Price of slaughter steers	.042	.267	.601
Price of utility cows	.058	.257	.587
Price of broilers	.036	.275	.562
Farm price of turkeys	.046	.245	.490
Farm price of eggs	.052	.353	.673
Farm price of milk	.025	.277	.591
Acreage planted of wheat	.030	.276	.546
Acreage planted of corn	.011	.177	.389
Acreage planted of barley	.037	.350	.728
Acreage planted of sorghum	.085	.542	1.770
Acreage planted of oats	.035	.412	.845
Acreage planted of soybeans	.042	.449	.943
Acreage planted of cotton	.047	.252	.499
Farm price of wheat	.062	.230	.456
Farm price of corn	.063	.300	.637
Farm price of barley	.083	.320	.804
Farm price of sorghum	.082	.421	.853
Farm price of oats	.036	.184	.366
Farm price of soybeans	.060	.258	.595
Farm price of cotton	.062	.302	.679
Total cash receipts from farm marketings	.016	.126	.259
Total farm production expenses	.010	.087	.181
Net farm income	.077	.279	.584
Consumer price index, all food	.019	.205	.490
Consumer price index, all items	.003	.036	.073

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